ESCUELA DE GOBIERNO

DOCUMENTOS DE TRABAJO 2026/08

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Abril 2026

Documentos de trabajo: https://bit.ly/2REorES

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Reserves as Insurance: International Buffers and Inward FDI in Emerging Markets

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Abstract

Emerging markets have accumulated large reserve buffers, but whether these buffers causally affect inward foreign direct investment (FDI) remains an open question. Using an unbalanced panel of emerging market economies over 2001–2020, we estimate two-way fixed-effects models of net inward FDI inflows with a rich set of lagged controls. We address the endogeneity of reserve accumulation by instrumenting lagged reserves with the two-year-lagged log of each country's commodity import price index—a source of balance-of-payments pressure orthogonal to export-driven profitability shocks, conditional on a country-specific commodity export price index. The IV estimates imply that a 10% increase in reserves raises FDI inflows by about 18.5% (an elasticity of 1.85), more than four times the fixed-effects OLS estimate of 0.4. The effect is amplified during global stress episodes: the IV elasticity is 1.85 in the full sample but falls to 1.34 when crisis years (2008–2009, 2020) are excluded, consistent with reserves functioning as insurance that matters most when downside risks are salient.

Keywords: international reserves; foreign direct investment; emerging markets; commodity prices; instrumental variables.

JEL codes: F21; F31; F32; E58.

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1 Introduction

Over the past two decades, emerging market economies have accumulated international reserves on an unprecedented scale. The classic motive for holding reserves is precautionary: reserves provide liquidity insurance against sudden stops and balance-of-payments crises (Obstfeld et al., 2010; Jeanne and Rancière, 2011; Durdu et al., 2009). Large reserve buffers may also matter outside crisis episodes by shaping investors' beliefs about policy credibility and the likelihood of disruptive adjustments when external conditions worsen. These considerations are especially relevant for foreign direct investment (FDI), a relatively stable source of external finance (Albuquerque, 2003; Forbes and Warnock, 2012) and a conduit for technology transfer. Despite the centrality of both reserves and FDI in emerging-market policy debates, we know relatively little about whether reserve accumulation causally affects inward FDI.

Theoretically, reserves can enhance macro-financial credibility and reduce perceived tail risks—sharp devaluations, capital controls, or abrupt policy reversals—that deter long-horizon investment (Julio and Yook, 2016; Busse and Hefeker, 2007; Azzimonti, 2018). Isolating this causal effect is challenging, however, because reserve accumulation is endogenous: governments often build buffers precisely when external vulnerability and unobserved risk factors worsen—conditions that independently depress FDI. As a result, both the sign and magnitude of the reserves–FDI relationship are theoretically ambiguous, and existing studies relying on correlations or structural models leave the causal question open. This paper fills that gap using an instrumental-variables strategy.

We estimate the effect of international reserves on inward FDI in emerging markets using an unbalanced panel covering 2001–2020. Our outcome is net FDI inflows as a share of GDP, and our main regressor is the stock of international reserves (alternatively scaled by GDP), entered with a one-year lag. We begin with a two-way fixed-effects specification with a rich set of lagged macroeconomic, institutional, and external-vulnerability controls. To address the endogeneity of reserve accumulation, we develop an instrumental variables strategy based on the lagged log level of each country's commodity import price index.

Specifically, we construct a country-specific import commodity price index from the IMF's Commodity Terms of Trade database using fixed historical import weights, and instrument for reserve holdings using lagged values of this index. The identifying idea is that changes in import commodity prices generate balance-of-payments pressures that affect reserve dynamics, while—conditional on country and year fixed effects, a comprehensive set of lagged controls, and a country-specific commodity export price index that absorbs export-driven profitability and revenue shocks—these import-price movements are plausibly orthogonal to the unobserved determinants of FDI at the country-year level.

In baseline fixed-effects estimates, higher lagged reserves are positively associated with FDI inflows, with an elasticity of 0.4. Instrumenting reserves with the two-year-lagged log of the import price index yields an IV elasticity of 1.85, more than four times the OLS estimate: a 10 percent increase in reserves raises FDI inflows by roughly 18.5 percent. The gap between OLS and IV estimates is consistent with downward bias in OLS arising from "negative selection" in reserve accumulation (countries tend to build buffers when unobserved risk is elevated), and may also reflect attenuation from measurement error in reserves. As in any single-instrument 2SLS design, this estimate should be interpreted as a local average treatment effect (LATE): it captures the causal effect of reserves on FDI for country-years in which reserve holdings respond to external shocks to the cost of the commodity import basket—rather than the effect of purely discretionary reserve accumulation. These complier episodes correspond to periods of balance-of-payments pressure driven by global commodity price fluctuations.

The findings are robust across reserve normalizations, alternative transformations of FDI that retain observations with zero or negative inflows, and standard treatments of outliers. We also document heterogeneity across global conditions. The IV elasticity falls from 1.85 to 1.34 when crisis years (2008–2009 and 2020) are excluded, consistent with an insurance-and-credibility channel in which reserves matter most during episodes of heightened global stress such as the global financial crisis and the COVID-19 shock, when downside risks are salient and external finance is more fragile (Frankel and Saravelos, 2012; Dominguez et al., 2012; Bussière et al., 2015).

The mechanism we emphasize is grounded in macro-financial risk perceptions. Higher international reserves signal external solvency and policy space, strengthening the government's capacity to smooth external shocks and reducing the likelihood of disruptive adjustment in bad states. This insurance-and-credibility channel implies that countries with larger reserve buffers should be more attractive destinations for FDI, particularly during global stress episodes when investors' sensitivity to downside risk is heightened—consistent with prior evidence on the importance of political and macroeconomic stability for FDI (Busse and Hefeker, 2007; Julio and Yook, 2016; Azzimonti, 2018) and with models of investment under uncertainty.[1] Section 6 provides suggestive evidence consistent with this channel.

This paper contributes in three ways. First, it provides cross-country causal evidence that reserve accumulation increases inward FDI in emerging markets. Second, by exploiting import commodity price variation rather than export-price shocks, our instrument isolates reserve variation driven by external financing pressures rather than export profitability. Third, the results bridge the reserves-as-insurance literature and the FDI determinants literature by highlighting a macro-financial channel through which policy buffers can support long-horizon investment.

The remainder of the paper is organized as follows. Section 2 reviews related literatures. Section 3 describes the data and variable construction. Section 4 presents the empirical framework and identification strategy. Section 5 reports the main results, robustness checks, and placebo timing tests. Section 6 provides suggestive evidence on mechanisms. Section 7 concludes.

[1] See Dixit and Pindyck (1994) on investment under uncertainty and Baker et al. (2016) on measuring economic policy uncertainty.

2 Related Literature

This paper connects two substantive literatures—the determinants of inward foreign direct investment (FDI) and the macro-financial role of international reserves in emerging markets—and draws methodologically on recent work on exposure-based designs that combine predetermined country weights with common shocks.

A large empirical literature studies the cross-country determinants of FDI, emphasizing market size, growth prospects, trade openness, institutions, and macroeconomic stability (Blonigen, 2005; Blonigen and Piger, 2014). Within this broad set of determinants, risk and uncertainty are central when projects are large, long-lived, and partially irreversible. Under imperfect capital markets, exchange rates also matter: Froot and Stein (1991) find that a weaker domestic currency is associated with higher FDI inflows.

More generally, irreversible-investment logic implies that uncertainty can delay entry. Using national elections as uncertainty shocks, Julio and Yook (2016) show that U.S. FDI flows to destination countries fall sharply in the quarter before elections; in their baseline, the flow rate declines by about 12% relative to non-election periods. Related evidence highlights the role of political risk and institutions for FDI (Busse and Hefeker, 2007; Alfaro et al., 2008), while Büthe and Milner (2014) emphasize credible commitment devices in international agreements.

This evidence motivates our central channel: if reserve buffers reduce perceived tail risk and the likelihood of disruptive policy responses, they can increase the attractiveness of long-horizon FDI, especially during global stress episodes. Models in which political instability raises expropriation risk support this mechanism (Azzimonti, 2018).

A second literature studies reserve accumulation and its macro-financial effects in emerging markets. Obstfeld et al. (2010) show that a one-standard-deviation rise in financial openness raises the predicted reserve-to-GDP ratio by about 0.16 log points (roughly 17%). Jeanne and Rancière (2011) and Durdu et al. (2009) similarly show that sudden-stop risk can imply higher precautionary reserve demand. Aizenman and Lee (2007) test precautionary versus mercantilist motives empirically, finding strong support for the former; Ghosh et al. (2017) show these motives have shifted over time as emerging markets became more financially integrated.

Related theory links reserves to sovereign financing risk: larger buffers reduce vulnerability to rollover crises and sudden stops (Caballero and Panageas, 2008; Bianchi et al., 2018). Empirical work also links reserve adequacy to better outcomes in stress episodes (Frankel and Saravelos, 2012; Bussière et al., 2015; Dominguez et al., 2012) and to broader gross capital-flow dynamics under stress: Alberola et al. (2016) show that reserve buffers dampen gross capital flow retrenchment during stress episodes, and that this buffering effect is stronger when pre-shock reserve levels are higher—consistent with a literature documenting that external (push) factors are dominant drivers of capital flow cycles to emerging markets (Calvo et al., 1993). Overall, reserves may shape private beliefs about external solvency and policy space beyond their mechanical liquidity role.

Closer to our question, Qian and Steiner (2014) study reserves and external equity composition, reporting that a 1% increase in reserves is associated with a 0.77% increase in the ratio of *portfolio equity investment* to FDI—suggesting that reserves may expand all forms of external equity, with portfolio equity responding more elastically than FDI. Complementary firm-level evidence in Tong and Wei (2021) shows that higher reserves-to-GDP are associated with higher corporate leverage, with stronger effects in sectors more exposed to uncertainty. Broto et al. (2011) find that a larger stock of reserves is associated with lower volatility of FDI inflows specifically—an effect not found uniformly across other capital flow types—suggesting that reserves stabilize not only the level but also the variability of inward direct investment.

Recent firm-level evidence by Aizenman et al. (2025) shows that active reserve management increases corporate investment in EMEs, with country credit spreads as the key transmission channel—a finding that complements our macro-level results and corroborates the sovereign risk mechanism we study in Section 6.

Yet credible cross-country causal evidence on reserves and the level of inward FDI remains limited, because reserve accumulation is endogenous to risk, policy credibility, and external conditions. We address this gap with an IV strategy for net inward FDI inflows.

Methodologically, our instrument uses the IMF Commodity Terms of Trade database (Gruss and Kebhaj, 2019): predetermined country-specific import weights interacted with global commodity price movements. This design belongs to the class of exposure-based instruments that combine predetermined country weights with common global shocks, so we draw on recent identification and inference results (Adão et al., 2019; Goldsmith-Pinkham et al., 2020; Borusyak et al., 2022). Because commodity-price dynamics and reserve management are closely linked—reserves buffer the exchange rate impact of terms-of-trade shocks in emerging markets (Aizenman and Riera-Crichton, 2008)—we condition on country-specific export price indices and rich lagged controls to isolate reserve variation driven by import-cost pressures rather than export-profitability shocks.

3 Data

Our empirical analysis uses an unbalanced panel of emerging market economies over the period 2001–2020. The estimation sample is determined by the joint availability of data on inward FDI inflows, international reserves, and the baseline set of macroeconomic and institutional controls. Appendix Table XII reports the country–year coverage of the baseline IV estimation sample.

The panel is unbalanced primarily because some variables are unavailable for certain countries and years, most notably sovereign spreads and institutional indicators. Despite this limitation, the data feature within-country time variation spanning up to 20 years per country (365 country-year observations across 27 countries), which is central to identification in the fixed-effects framework.

The dependent variable is net foreign direct investment inflows as a share of GDP. We obtain FDI inflows from the World Development Indicators (WDI), using the standard balance-of-payments definition that includes net equity flows and reinvested earnings. In the baseline specification, the dependent variable is the logarithm of FDI/GDP. Because FDI inflows can be zero or negative in some country-years (reflecting net divestment), the log transformation is defined only for strictly positive values. We therefore restrict the baseline log specification to observations with $FDI/GDP > 0$, which excludes about 5% of the available sample. In robustness checks, we use transformations defined on the full real line—including the inverse hyperbolic sine transformation and shifted-log specifications—to retain observations with non-positive FDI inflows.

The main explanatory variable is the stock of international reserves, measured according to the IMF definition (foreign exchange assets, monetary gold, Special Drawing Rights, and the reserve position at the IMF). Reserve data are taken primarily from the WDI and complemented with IMF sources when needed. In the baseline analysis, reserves enter in logarithms and are lagged one year. We consider two reserve measures: (i) total reserves in levels and (ii) reserves as a share of GDP.

Appendix Table X reports the full list of variables, together with definitions and sources. To isolate the effect of reserves on FDI inflows, the baseline specifications include a set of one-year lagged controls capturing macroeconomic conditions (e.g., GDP per capita and inflation), external vulnerability, institutional and policy conditions, and structural characteristics. Unless explicitly stated otherwise, the baseline regressions in the paper use this same control set.

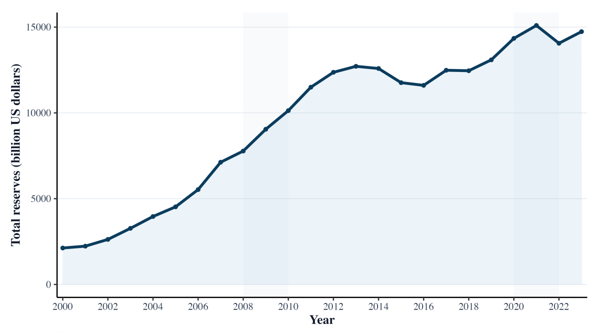
Commodity price movements provide a key source of external variation for emerging markets, so we draw on two country-specific commodity-price indices from the IMF Commodity Terms of Trade (CTOT) database. We control for the Export Price Index, which weights international commodity prices by each country's export shares. This index captures commodity-driven export revenue and profitability shocks that may directly affect income, competitiveness, and investment incentives (including FDI in commodity-related sectors). We also use an import commodity price index that tracks changes in the cost of the imported commodity basket using fixed historical import weights (average shares over 1980–2015) and deflated by a manufactured-goods price index. The import price index is conceptually distinct from the export index—capturing shocks to import costs rather than export profitability—and is used exclusively as an instrument in the IV strategy described in Section 4.

The institutional variable (*Corruption*) contains missing observations for 2001, which we impute using the country-specific average of adjacent years (2000 and 2002). Appendix Table XXI shows that excluding these imputed values yields nearly identical results: coefficient magnitudes change only marginally, and statistical significance is unaffected. This suggests that the imputation does not materially influence the baseline findings.

3.1 Overview of Reserve and FDI Dynamics

Figure I documents the evolution of international reserves and FDI inflows over the sample, highlighting the sources of variation that motivate the empirical strategy.

(a) International Reserves: Annual Sum (2000–2023)



(b) Reserves-to-GDP and FDI-to-GDP (2000–2023)

Figure I: International reserves and FDI dynamics, 2000–2023.

Panel (a) plots the annual sum of total international reserves (including gold) across countries with available data in each year, expressed in current U.S. dollars (billions). In other words, each point is the cross-country aggregate reserve stock for that year, not a mean or median country value. The series shows that the aggregate reserve stock grew roughly seven-fold in nominal terms, from roughly \$2,000 billion in 2000 to over \$14,000 billion by 2023—an increase that exceeds cumulative U.S. dollar inflation of approximately 82%, reflecting a genuine strengthening of external buffers rather than mere nominal revaluation. Reserve growth was most rapid between 2003 and 2010, slowed during the global financial crisis (first shaded area) as several countries drew down buffers to stabilize external conditions, and resumed through the 2010s. A further inflection is visible during the COVID-19 pandemic (second shaded area).

Panel (b) overlays annual means of Reserves-to-GDP and FDI-to-GDP in a single dual-axis figure and shows different dynamics. Reserves-to-GDP follows a secular upward pattern over the sample, while FDI-to-GDP exhibits more cyclical variation, including a decline around global stress episodes and only partial recovery thereafter. Taken together, the panels show that reserves and FDI do not move in lockstep, underscoring the need to account for confounders when estimating the impact of reserves on FDI inflows.

Two features of the data are central to the empirical strategy. First, aggregate reserve levels grow steadily across subperiods in Panel (a), while in Panel (b) the reserves-to-GDP ratio follows a secular upward trend and FDI-to-GDP is more sensitive to global cycles. Second, the two series show distinct time profiles rather than tight comovement. Together, these patterns motivate the use of exogenous variation in reserve holdings to isolate causal effects, as described in Section 4.

4 Empirical Framework

4.1 Baseline specification

We estimate the effect of international reserves on inward FDI using a panel specification with country and year fixed effects. Let i index countries and t years. The baseline regression is

$$y_{i,t} = \beta\, r_{i,t-1} + \gamma' X_{i,t-1} + \phi\, \log E_{i,t-2} + \mu_i + \tau_t + \varepsilon_{i,t}, \tag{1}$$

where $y_{i,t} \equiv \log(FDI_{i,t}/GDP_{i,t})$ is the logarithm of net FDI inflows as a share of GDP, $r_{i,t-1}$ is a lagged measure of reserves, $X_{i,t-1}$ is a vector of one-period-lagged controls, $E_{i,t}$ is the commodity export price index (IMF CTOT), and μ_i and τ_t are country and year fixed effects. The export price index enters at $t-2$ rather than $t-1$ so that it is contemporaneous with the excluded instrument $z_{i,t} \equiv \log M_{i,t-2}$, ensuring that the exclusion restriction is conditioned on export-price-driven commodity shocks measured at the same horizon. Since global commodity prices comove, conditioning on the export price index at the same lag as the instrument removes the component of import-price variation that reflects broad commodity price cycles common to both baskets, leaving residual variation in import costs more plausibly orthogonal to FDI determinants.

The regressor $r_{i,t-1}$ is defined as the logarithm of the stock of international reserves. In alternative specifications, we replace $r_{i,t-1}$ with the log of the reserves-to-GDP ratio, as described in Section 3. The coefficient β therefore has an elasticity interpretation. Equation (1) is estimated with standard errors two-way clustered by country and year.

4.2 Instrumental variables strategy

A central concern is that reserve accumulation may be endogenous to unobserved risk, policy credibility, or investor sentiment that also affects FDI. To identify the causal impact of reserves on inward FDI, we implement a two-stage least squares (2SLS) strategy that instruments lagged reserves with the two-year lag of the log import commodity price index for each country ($z_{i,t} \equiv \log M_{i,t-2}$). The identifying idea is that import commodity price shocks plausibly shift reserves through balance-of-payments pressures and reserve management, while—conditional on fixed effects and the full control set (including the commodity export price index)—they are unlikely to affect inward FDI through other direct channels.

4.2.1 Instrument construction

We use a country-specific import commodity price index constructed from the IMF Commodity Terms of Trade (CTOT) database (see Section 3). Let $M_{i,t}$ denote the import price index for country i in year t (normalized to 2012 = 100). The index is constructed by cumulating annual log changes in commodity prices using fixed historical import weights:

$$\log M_{i,t} = \log M_{i,t-1} + \sum_{j=1}^{J} w_{i,j}\, \Delta \log P_{j,t}, \qquad w_{i,j} = \frac{\bar{m}_{i,j}}{\sum_{k=1}^{J} \bar{m}_{i,k}}, \tag{2}$$

where $P_{j,t}$ is the (real) international price of commodity j and $\bar{m}_{i,j}$ is the average import value of commodity j in country i over the reference period used to compute fixed weights. The instrument combines predetermined country-specific import shares with global commodity price movements. We therefore follow the recent econometric literature that clarifies identification and inference in exposure-based designs (Adão et al., 2019; Goldsmith-Pinkham et al., 2020; Borusyak et al., 2022). In the baseline IV specification, the instrument is the two-year lag of the log index,

$$z_{i,t} \equiv \log M_{i,t-2}. \tag{3}$$

Because $M_{i,t}$ is constructed by cumulating log price changes, $\log M_{i,t-2}$ summarizes lagged import-price movements facing each country, scaled by predetermined import weights. The two-period lag allows time for import price shocks to transmit to balance-of-payments pressures and reserve accumulation decisions by central banks, while maintaining sufficient instrument strength.

4.2.2 Two-stage least squares

The first stage relates lagged reserves to the lagged import price index:

$$r_{i,t-1} = \pi\, z_{i,t} + \delta' X_{i,t-1} + \psi\, \log E_{i,t-2} + \mu_i + \tau_t + \nu_{i,t}. \tag{4}$$

The relevance of commodity-price shocks for reserve accumulation is consistent with empirical evidence that external shocks, including commodity price movements, are an important driver of reserve dynamics in emerging markets (Aizenman and Riera-Crichton, 2008).

The second stage substitutes the fitted values from (4) into the outcome equation:

$$y_{i,t} = \beta \, \widehat{r}_{i,t-1} + \gamma' X_{i,t-1} + \phi \, \log E_{i,t-2} + \mu_i + \tau_t + \varepsilon_{i,t}. \tag{5}$$

Both stages include the same fixed effects and control set as in the baseline specification. Standard errors are two-way clustered by country and year.

4.2.3 Identification and diagnostics

Our IV strategy exploits cross-country variation in import commodity prices to identify the causal effect of reserve holdings on inward FDI. The identifying assumption is that, conditional on country and year fixed effects and a rich set of lagged controls $X_{i,t-1}$, movements in the import commodity price index affect FDI primarily through their impact on reserve accumulation rather than through other direct channels.

A potential threat to identification is that import price shocks may influence the investment environment directly, for instance through persistent effects on production costs, inflation, or aggregate activity. These concerns mirror the identifying-assumption issues emphasized in exposure-based designs, where validity requires that exposure shares interact with shocks that are as-good-as-random with respect to unobserved outcomes once the relevant controls are included (Goldsmith-Pinkham et al., 2020; Borusyak et al., 2022).

Several features of the empirical design mitigate these concerns. First, the instrument is constructed from *import* commodity prices rather than export or net export prices, and thus captures changes in the cost of the import basket rather than shifts in export profitability, since export-price shocks may directly affect FDI by changing investment incentives in commodity sectors, not merely through their effect on reserves. A potential concern is that higher import costs may depress FDI by raising production costs. However, conditional on inflation and GDP growth controls, this channel is unlikely to confound our estimates. Second, the specification controls for a country-specific commodity Export Price Index, which absorbs export-price-driven revenue and profitability shocks that could otherwise affect FDI directly (for example by changing investment incentives in commodity-related sectors). Because global commodity prices often comove, conditioning on export prices also removes the component of import-price variation that reflects broad commodity booms.

Conditional on the Export Price Index and lagged levels of second-stage controls such as inflation and growth, residual variation in import prices more plausibly reflects import-cost pressures and balance-of-payments dynamics that operate through reserves. Related IV strategies use shocks to commodity import prices to isolate externally driven macroeconomic variation: Garcia-Macia (2023), for instance, instruments inflation with a country-specific import commodity price index.

Third, the instrument is lagged two periods, introducing temporal separation between import price shocks and contemporaneous FDI inflows. This timing is consistent with the institutional and economic mechanisms governing reserve accumulation, which typically respond to external price movements with delay due to trade invoicing, balance-of-payments adjustment, and policy implementation. While alternative lag choices yield qualitatively similar conclusions, the two-period lag delivers the highest first-stage F-statistic ($F = 19.8$, versus $F = 16.4$ for $L = 3$ and $F = 3.2$ for $L = 1$); Appendix Table XVI reports results for $L \in \{1, 2, 3\}$.

We complement the identification strategy with a comprehensive set of diagnostics. In the main IV results in Section 5, we report first-stage strength measures and endogeneity diagnostics, and we complement conventional inference with additional robustness checks. Additional robustness checks are reported in Appendix F.

5 Results

This section reports the main empirical results on the relationship between international reserves and inward foreign direct investment (FDI). We begin with two-way fixed-effects OLS estimates of equation (1) and then present instrumental variables (2SLS) estimates based on the identification strategy described in Section 4. We next assess robustness to functional-form choices and sample restrictions, and finally study dynamic responses using LP-IV. Unless otherwise noted, all regressions include country and year fixed effects, the full set of lagged controls listed in Section 3, and standard errors two-way clustered by country and year.

5.1 Baseline fixed-effects estimates

Table I reports baseline within-country estimates of the effect of reserves on inward FDI inflows. Across specifications, the coefficient on lagged reserves is positive and statistically significant, and it is stable across alternative reserve measures (levels and reserves scaled by GDP).

Table I: Baseline fixed-effects estimates: international reserves and FDI inflows

Dependent Variable:	(1)	(2)
$log(Res)_{t-1}$	0.419***	
	(0.124)	
$log(Res/GDP)_{t-1}$		0.431***
		(0.109)
Observations	365	365
Adj. R^2	0.501	0.502

Fixed effects and Clustered (Country and Time). Standard-errors in parentheses

*Signif. Codes: *** 0.01, ** 0.05, * 0.1*

The magnitude is economically meaningful. In the baseline specification, the estimated elasticity of 0.419 implies that a 10% increase in reserves is associated with roughly a 4.19% increase in FDI inflows. At the 2023 average values for the sample countries (Reserves/GDP = 22.45%, FDI/GDP = 3.5%),[2] a 10% increase in reserves to GDP (22.45% to 24.7%) would mean FDI inflows increasing from 3.5% to approximately 3.65% of GDP.

5.2 Instrumental variables estimates

We next address the potential endogeneity of reserve accumulation. Even with extensive controls and country and year fixed effects, reserve holdings may respond to unobserved shifts in country risk, external financing conditions, or investor sentiment that also affect inward FDI. In that case, fixed-effects OLS may be biased, for instance if countries accumulate reserves precisely when macroeconomic or political risks are elevated and foreign investors retrench. To isolate plausibly exogenous movements in reserves, we instrument reserve holdings with the two-year-lagged log of each country's commodity import price index ($z_{i,t} \equiv \log M_{i,t-2}$).

Table II reports the first- and second-stage estimates using the baseline reserve measure and reserves scaled by GDP. The first stage confirms that import commodity prices predict subsequent reserve accumulation, with a first-stage F-statistic of 19.8 and 19.5, respectively. Importantly, the results are not driven by sample composition in the unbalanced panel. In Appendix B, we re-estimate the IV specification on a restricted subpanel that requires countries to have at least 12 consecutive years of data, thereby addressing concerns about selection through panel imbalance. The first-stage relevance and second-stage coefficients remain positive and statistically significant, reinforcing the robustness of the causal interpretation.

The second-stage estimates imply a substantially larger effect of reserves on FDI than suggested by fixed-effects OLS. The estimated IV elasticity is 1.850, implying that a 10% increase in reserves raises FDI inflows by about 18.5%. At the 2023 average values for the sample countries (Reserves/GDP = 22.45%, FDI/GDP = 3.5%), a 10% increase in reserves to GDP (22.45% to 24.7%) would mean FDI inflows increasing from 3.5% to approximately

[2]These values are drawn from the World Development Indicators for 2023 and lie outside the estimation sample (2001–2020); they are used here solely for illustrative purposes.

4.18% of GDP—a gain of about 0.68 percentage points.

Table II: IV Estimation: Reserves and Reserves-to-GDP

	(1) FS	(2) SS	(3) FS	(4) SS
$\log(M_i)_{t-2}$	1.436***		1.436***	
	(0.386)		(0.420)	
$\widehat{\log(Res)}_{t-1}$		1.850***		
		(0.557)		
$\log(\widehat{Res/GDP})_{t-1}$				1.851***
				(0.510)
Controls	Yes	Yes	Yes	Yes
Country and Year FE	Yes	Yes	Yes	Yes
Observations	365	365	365	365
First-stage F-stat (excluded instruments)	19.8		19.5	
Wu–Hausman p-value		0.00946		0.00999
Anderson–Rubin p-value ($H_0 : \beta = 0$)		0.000136		0.000136

Notes: Columns (1)-(2) use reserves in levels; columns (3)-(4) use reserves-to-GDP. Standard errors are two-way clustered by country and year. The reported first-stage F-statistic corresponds to the excluded instrument in the first stage.

**** p<0.01, ** p<0.05, * p<0.10.*

The gap between OLS and IV estimates is consistent with downward bias in OLS due to negative selection, and potentially with attenuation from measurement error in reserves. In other words, countries that are less attractive destinations for FDI may be systematically accumulating more reserves as a precautionary buffer. Additional diagnostics and sensitivity checks (reported in the Appendix F) are consistent with this interpretation. As a further robustness check, Appendix Table XIX reports results from an overidentified specification that augments the baseline import-price instrument with lagged SDR allocations from the IMF; the estimated IV elasticity remains stable at 1.6–1.7, consistent with the baseline.

5.3 Robustness, heterogeneity, and placebo tests

This subsection assesses robustness along three dimensions: (i) alternative transformations of the dependent variable (including those that retain observations with zero or negative FDI inflows), (ii) alternative reserve measures, and (iii) heterogeneity across global conditions and targeted subsamples. OLS robustness estimates across these same transformations and reserve normalizations are reported in Appendix Table XIV.

5.3.1 Robustness: IV estimates across reserve measures

Table III reports IV estimates for (i) reserves in levels and (ii) reserves scaled by GDP across four transformations of FDI/GDP: log (baseline), IHS, shifted log, and levels, with winsorized variants for the latter two. The IV elasticities are consistently larger than the corresponding OLS estimates, consistent with downward bias in OLS from negative selection in reserve accumulation.

Table III: IV robustness across reserve measures

	Log-based		IHS-based		Levels	
Dep. var.: FDI/GDP	(1)	(2)	(3)	(4)	(5)	(6)
	Log	Shifted log	IHS	IHS (wins.)	Levels	Levels (wins.)
Panel A: $\widehat{\log(Res)}_{t-1}$						
$\widehat{\log(Res)}_{t-1}$	1.850***	1.836***	1.904***	1.759***	8.754**	6.004***
	(0.557)	(0.551)	(0.598)	(0.533)	(3.525)	(1.948)
Observations	365	365	372	372	372	372
Adj. R^2	0.306	0.309	0.414	0.456	0.135	0.402
Panel B: $\log(\widehat{Res/GDP})_{t-1}$						
$\log(\widehat{Res/GDP})_{t-1}$	1.851***	1.837***	1.936***	1.788***	8.900**	6.104***
	(0.510)	(0.504)	(0.545)	(0.484)	(3.369)	(1.815)
Observations	365	365	372	372	372	372
Adj. R^2	0.308	0.310	0.386	0.435	0.090	0.383

Significance: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Notes: Each column applies a different transformation to the dependent variable FDI/GDP. *Shifted log* uses $\log(y + \kappa)$, where κ shifts the distribution to retain non-positive observations. *IHS* applies the inverse hyperbolic sine transformation, $\sinh^{-1}(y)$, which approximates the log for large values while accommodating zeros and negative values. *IHS (wins.)* and *Levels (wins.)* winsorize the dependent variable at the 1st and 99th percentiles before applying the transformation.

5.3.2 Subsample checks: global stress episodes

Table IV reports estimates excluding major global stress years (2008–2009 and 2020). Excluding crisis years reduces the magnitude of the IV coefficient, but the estimated effect remains positive and statistically significant, consistent with a channel in which reserves matter most when global financial conditions deteriorate.

Because the specification uses lagged reserves and lagged controls, exclusion exercises can be defined more or less stringently. A conservative approach is to exclude not only the stress years themselves but also adjacent years for which lagged regressors would mechanically draw on stress-year values. Appendix F.1.1 reports targeted exclusion exercises that just drop the SDR allocation year (2009).

Table IV: Subsample checks: excluding global crisis years (2008–2009 and 2020)

Specification:	(1)	(2)	(3)	(4)
$\log(Res)_{t-1}$	0.337**			
	(0.134)			
$\log(Reserves/GDP)_{t-1}$			0.386***	
			(0.097)	
$\widehat{\log(Res)}_{t-1}$		1.340***		
		(0.371)		
$\log(\widehat{Reserves}/GDP)_{t-1}$				1.335***
				(0.276)
Observations	288	288	288	288
Adj. R^2	0.494	0.392	0.499	0.402

Country and year fixed effects; standard errors two-way clustered (country, year).
*Significance levels: *** p<0.01, ** p<0.05, * p<0.10*

Columns (1) and (3) are estimated using OLS. Columns (2) and (4) use the previously described IV estimator.

This crisis-state dependence is consistent with earlier cross-country evidence that reserves are particularly valuable during periods of global stress (Frankel and Saravelos, 2012; Dominguez et al., 2012; Bussière et al., 2015), and with evidence that the dampening effect of reserves on gross capital flow retrenchment is itself increasing in the pre-shock reserve buffer (Alberola et al., 2016)—a pattern directly consistent with our finding of larger reserve–FDI elasticities during the 2008–2009 crisis and COVID-19.[3]

5.3.3 Placebo timing tests

We assess whether the estimated reserve–FDI relationship could reflect anticipation effects (or reverse timing) by testing whether future reserve holdings predict current FDI, once we condition on fixed effects and the baseline control set. OLS placebo timing regressions are reported in Appendix D. Table V reports the IV placebo estimates for both reserve measures: Panel A uses reserve levels and Panel B uses reserves as a share of GDP. Columns (1)–(3) report *single-horizon* 2SLS estimates in which a single reserve term enters the second stage (at t, $t+1$, and $t+2$, respectively) and is instrumented using the same import-price instrument shifted in time (i.e., the import-price index at $t-1$, t, and $t+1$, respectively). Column (4) reports the joint placebo specification: we simultaneously instrument the reserve variable at t, $t+1$, and $t+2$ using the import-price measure at the corresponding three different dates.

[3]We note that FDI can also rise during liquidity crises through fire-sale dynamics, as foreign acquirers exploit distressed valuations (Aguiar and Gopinath, 2005). Our mechanism—reserves reducing perceived tail risk ex ante—is conceptually distinct: higher reserve buffers attract FDI by improving the macro-financial environment before a crisis materializes, not through post-crisis fire-sale acquisitions.

Table V: Placebo timing test: reserves and FDI

Specification:	(1)	(2)	(3)	(4)
Panel A: Reserve levels, log(*Res*)				
$\log(Res)_t$	2.621			-16.177
	(2.055)			(57.149)
$\log(Res)_{t+1}$		10.789		17.336
		(44.590)		(57.392)
$\log(Res)_{t+2}$			-1.767	-2.655
			(3.291)	(14.854)
Observations	365	365	365	365
Adj. R^2	0.164	-9.898	0.019	-12.435
Panel B: Reserves as % of GDP, log(*Res/GDP*)				
$\log(Res/GDP)_t$	2.639			-3.337
	(1.997)			(6.298)
$\log(Res/GDP)_{t+1}$		-124.659		4.144
		(5184.177)		(6.692)
$\log(Res/GDP)_{t+2}$			-4.562	0.869
			(17.660)	(3.082)
Observations	365	365	365	365
Adj. R^2	0.167	-1555.957	-1.653	-0.745

Country and year fixed effects; standard errors two-way clustered (country, year).

*Significance levels: *** p<0.01, ** p<0.05, * p<0.10*

Panel A uses reserve levels; Panel B uses reserves as a share of GDP.

In both panels, leads of reserves do not predict current FDI inflows, which strengthens the causal interpretation of the baseline results.

5.4 Dynamic effects of reserves on FDI: local projections with IV

The static 2SLS estimates reported above identify an average effect of lagged reserves on contemporaneous FDI inflows. For interpretation and policy relevance, it is also useful to characterize when this effect materializes and how long it persists. We therefore complement the baseline IV results with an instrumental-variables local-projections approach (Jordà, 2005; Stock and Watson, 2018), which traces the dynamic response of FDI to an exogenous increase in reserves without imposing a tightly parameterized dynamic structure.

LP-IV specification. For each horizon $h = 1, 2, \dots, H$, we estimate a separate 2SLS regression where the dependent variable is FDI at $t + h - 1$ and the endogenous regressor is reserves at $t - 1$, instrumented with the same lagged import commodity price index used in the main IV strategy. At $h = 1$ the specification is identical to the static IV. Specifically, for each h we estimate:

$$\textbf{(First stage)} \quad r_{i,t-1} = \pi\, z_{i,t} + \delta' X_{i,t-1} + \psi\, \log E_{i,t-2} + \mu_i + \tau_{t+h-1} + \nu_{i,t-1}, \tag{6}$$

$$\textbf{(Second stage)} \quad y_{i,t+h-1} = \beta_h\, \widehat{r}_{i,t-1} + \gamma_h' X_{i,t-1} + \phi\, \log E_{i,t-2} + \mu_i + \tau_{t+h-1} + \varepsilon_{i,t+h-1}, \tag{7}$$

where $y_{i,t+h-1}$ is a transformation of net FDI inflows as a share of GDP, $r_{i,t-1}$ is the log of reserves (or log reserves/GDP) lagged one year, $X_{i,t-1}$ is the baseline set of one-period-lagged controls, $E_{i,t-2}$ is the commodity export price index entering at $t - 2$, μ_i are country fixed effects, and τ_{t+h-1} are year fixed effects indexed to the outcome year. The instrument $z_{i,t} \equiv \log M_{i,t-2}$ is the same two-year lag of the import commodity price index used in the static IV.

We report three outcome transformations to assess robustness to outliers and to retain observations with zero or negative net inflows: (i) log(FDI/GDP) (defined only for positive inflows), (ii) IHS(FDI/GDP), and (iii) winsorized IHS(FDI/GDP).

Figure II summarizes the estimated impulse responses $\{\widehat{\beta}_h\}_{h=1}^{8}$. For the baseline outcome ($\log$ FDI/GDP) and reserve levels, the estimated elasticity is 1.850*** at $h = 1$, 1.530** at $h = 2$, 0.786** at $h = 3$, and 1.274** at $h = 4$; the corresponding estimates for reserves-to-GDP are 1.851***, 1.572**, 0.808**, and 1.306** (Table VI, columns 4–7). The positive and statistically significant responses at $h = 2$, $h = 3$ and $h = 4$ indicate that the effect is not merely transitory: reserve-induced improvements in perceived macro-financial safety propagate through an investment "pipeline" over multiple years, consistent with staggered project execution, approval processes, and gradual scaling of multinational activity.

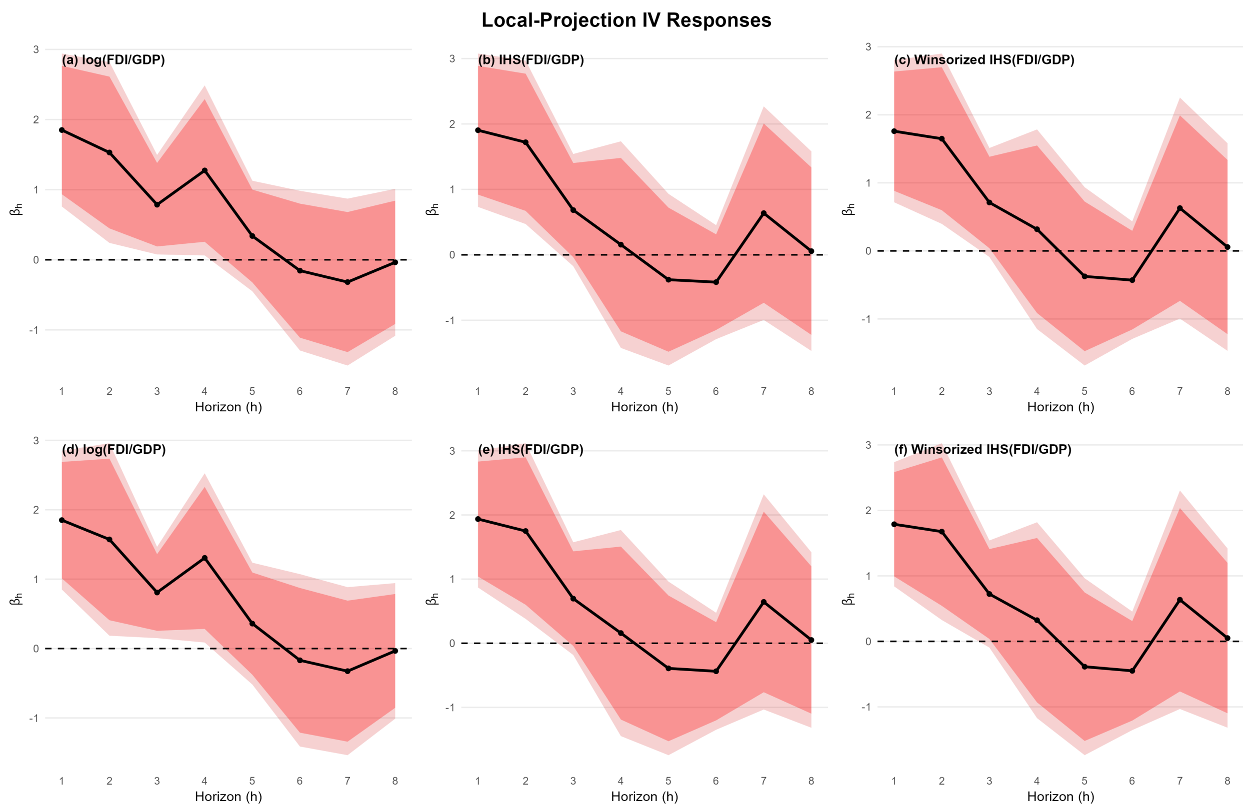


Figure II: **LP-IV: Dynamic response of inward FDI to reserves.** Notes: Panel-specific LP-IV estimates ($h = 1, \ldots, 8$) with 90/95% two-way clustered confidence bands.

Cumulative LP-IV responses are reported in Appendix Figure III. The cumulative elasticity stabilizes at approximately 5 from $h = 5$ onward—implying a cumulative FDI/GDP gain of about 50% from a 10% reserve increase—though this estimate should be read as an upper bound given widening confidence bands at longer horizons.

5.5 Summary of main findings

This section delivers three core results. Baseline OLS estimates (Table I) indicate a positive and statistically significant association between lagged reserves and inward FDI, with an elasticity of approximately 0.4. IV estimates (Table II) imply a substantially larger elasticity of 1.85—more than four times the OLS estimate—consistent with downward bias from negative selection. This positive effect is robust to alternative functional forms, reserve measures, and sample composition, with OLS robustness checks (Table XIV) and IV robustness checks (Table III) consistently yielding the same sign and relative ordering across both reserve normalizations (levels and reserves-to-GDP).

Finally, LP-IV impulse responses show that the effects are concentrated at short horizons but can persist beyond the impact horizon, with statistically significant responses through $h = 4$ in the baseline specification (Figure II). Cumulative responses are correspondingly front-loaded: the cumulative elasticity stabilizes at approximately 5 after five periods (Appendix Figure III), so that a 10% reserve increase implies a cumulative FDI-to-GDP gain of about 50%. Overall, the evidence supports the paper's central conclusion: exogenous increases in international reserves raise inward FDI, with a dynamic profile concentrated at $h = 1$–4 and detectable over a short-to-intermediate window.

Placebo timing tests (Table V) confirm that the results are not driven by anticipation effects or reverse timing: future reserve holdings do not predict current FDI inflows once we condition on fixed effects, the baseline controls, and the contemporaneous reserve level.

Table VI collects the key estimates. Excluding crisis years attenuates the IV estimate from 1.85 to 1.34, confirming that the effect is amplified during episodes of elevated global stress.

Table VI: Summary of key elasticity estimates

	(1) OLS	(2) IV	(3) IV (excl. crisis)	(4) LP-IV ($h = 1$)	(5) LP-IV ($h = 2$)	(6) LP-IV ($h = 3$)	(7) LP-IV ($h = 4$)
$\log(\text{Res})_{t-1}$	0.419***	1.850***	1.340***	1.850***	1.530**	0.786**	1.274**
	(0.124)	(0.557)	(0.371)	(0.557)	(0.658)	(0.362)	(0.618)
$\log(\text{Res/GDP})_{t-1}$	0.431***	1.851***	1.335***	1.851***	1.572**	0.808**	1.306**
	(0.109)	(0.510)	(0.276)	(0.510)	(0.708)	(0.336)	(0.622)
Observations	365	365	288	365	364	363	362

Notes: Each cell reports the coefficient on the reserve measure indicated in the row. Dependent variable is $\log(\text{FDI/GDP})_t$ in columns (1)–(3) and $\log(\text{FDI/GDP})_{t+h-1}$ in columns (4)–(7). Column (1) is OLS; columns (2)–(3) instrument the reserve measure with $\log M_{i,t-2}$. Column (3) excludes crisis years 2008–2009 and 2020. Columns (4)–(7) report LP-IV horizon-specific elasticities; standard errors recovered from 95% confidence intervals. $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.10$.

6 Suggestive Evidence on Mechanisms

This section probes two risk-reduction channels—lower sovereign risk premia and a reduced probability of policy tightening—through which reserves may affect inward FDI. We construct empirical proxies for each and ask whether conditioning on them attenuates the estimated reserve–FDI relationship. The evidence should be read as consistency checks rather than causal decompositions.

6.1 Sovereign risk channel

Following Gómez et al. (2025), we regress $\log(\text{Spread}_{i,t})$ on lagged global and domestic fundamentals $Z_{i,t-1}$ and on changes in reserves attributed to three financing sources:

$$\log(\text{Spread}_{i,t}) = \Gamma' Z_{i,t-1} + \theta_{\text{priv}}\,\Delta R^{\text{priv}}_{i,t} + \theta_{\text{pub}}\,\Delta R^{\text{pub}}_{i,t} + \theta_{\text{IMF}}\,\Delta R^{\text{IMF}}_{i,t} + \mu_i + \tau_t + u_{i,t}, \quad (8)$$

where ΔR^{priv}, ΔR^{pub}, and ΔR^{IMF} denote reserve changes attributable to private flows, public flows, and IMF lending, respectively.

Table VII is consistent with a sovereign-risk channel. Reserve accumulation driven by private and public external flows is associated with lower sovereign spreads, as reflected in negative and statistically significant coefficients. This pattern is in line with theoretical work in which reserves provide insurance against rollover risk and sudden stops, thereby reducing sovereign risk in adverse states (Caballero and Panageas, 2008; Bianchi et al., 2018). To the extent that sovereign spreads summarize external financing conditions and macro-financial tail risk, these results are also plausibly relevant for multinational firms' location decisions. By contrast, reserve changes linked to IMF lending are associated with higher spreads, consistent with IMF-supported episodes tending to occur in periods of elevated stress. We use the fitted values from equation (8), $\widehat{\log(\text{Spread})}_{i,t-1}$, as an empirical proxy for sovereign risk in the augmented regressions of Section 6.3.

Table VII: Sovereign spread regressions: sources of reserve accumulation

Variable	Coefficient
$\Delta R^{private}$	-0.768***
	(0.176)
ΔR^{public}	-0.596**
	(0.217)
ΔR^{IMF}	1.685**
	(0.809)
Observations	896
Adj. R^2	0.775

Country and year fixed effects; standard errors two-way clustered (country, year).
*Significance: *** p<0.01, ** p<0.05, * p<0.10.*

6.2 Predicted probability of policy tightening

We construct a proxy for the risk that a country experiences an abrupt deterioration in policy/institutional openness—a dimension of macro-policy risk that is plausibly salient for multinational investors and therefore for inward FDI. This motivation is consistent with evidence that political and policy uncertainty and institutional risk are first-order determinants of FDI, and with work emphasizing the role of credible commitment devices for sustaining foreign investment (Julio and Yook, 2016; Busse and Hefeker, 2007; Büthe and Milner, 2014; Azzimonti, 2018). We define the binary tightening indicator:

$$\texttt{tighten\_policy}_{i,t} = \mathbb{1}\Big\{\texttt{InstitutionAvg}_{i,t} < 0.99 \cdot \texttt{InstitutionAvg}_{i,t-1}\Big\}, \tag{9}$$

where $\texttt{InstitutionAvg}_{i,t}$ averages "Freedom to trade internationally" and "Regulation" from the Economic Freedom of the World dataset. The threshold requires a decline of more than 1% to flag a tightening episode.[4] We then estimate a panel logit model with country and year fixed effects—entering reserves in first differences—to obtain fitted tightening probabilities $\widehat{p}_{i,t}$.

Table VIII shows that changes in reserves enter with a negative and statistically significant coefficient in both specifications. Interpreted in log-odds units, this implies that larger reserve buffers are associated with a lower predicted probability of a tightening episode, conditional on country and year fixed effects and the full set of lagged controls. We use $\widehat{p}_{i,t-1}$ as an additional control in augmented versions of the main FDI regressions.

[4] Results with $\kappa = 1$ are reported in Appendix Table XX.

Table VIII: Policy-tightening logit: reserve coefficients

	(1)	(2)
$\Delta \log(Res/GDP)_t$	-1.964*	–
	(1.056)	
$\Delta \log(Res)_t$	–	-2.038**
		(0.944)
Country and Year FE	Yes	Yes
Controls	Yes	Yes
Observations	372	372

Significance: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

6.3 Augmented FDI regressions

We add $\log(\widehat{\text{Spread}})_{i,t-1}$ and $\widehat{p}_{i,t-1}$ as additional controls to the baseline IV specification. Table IX shows that the reserve coefficient attenuates once these proxies are included.

Taken together, the results in this section are consistent with the insurance-and-credibility mechanism proposed in this paper. The estimated reserve–FDI elasticity attenuates when we condition on proxies for sovereign risk and policy tightening risk, in the direction one would expect if reserves partly operate by reducing perceived tail risk. The reserve coefficient falls from 1.85 to 1.58. We do not claim to have decomposed the causal effect of reserves into its constituent channels — the proxies are themselves endogenous and generated in preliminary steps, and the augmented specifications should not be interpreted as a mediation analysis. What the evidence does support is that the macro-financial risk environment is a plausible and empirically relevant pathway, consistent with the broader finding that reserve buffers matter most precisely when downside risks are most salient.

Table IX: Augmented IV Regressions: FDI with Sovereign Risk and Policy Tightening Proxies

Dep. var.: FDI/GDP	(1)	(2)
	$\widehat{\log(Res)}_{t-1}$	$\log(\widehat{Res/GDP})_{t-1}$
$\widehat{\log(Res)}_{t-1}$	1.581**	
	(0.724)	
$\log(\widehat{Res/GDP})_{t-1}$		1.523**
		(0.687)
$\widehat{\log(\text{Spread})}_{t-1}$	0.345	0.264
	(0.286)	(0.245)
$Prob(\widehat{\text{Policy Tight.}})_{i,t-1}$	-0.344	-0.542*
	(0.270)	(0.307)
Observations	331	331
Adj. R^2	0.437	0.439

Significance: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Notes: $\widehat{\log(\text{Spread})}_{i,t-1}$ and $Prob(\widehat{\text{Str. Cap. Cont.}})_{i,t-1}$ are fitted values from auxiliary regressions as described in Section 6.

7 Conclusion

This paper asks whether international reserves causally affect inward foreign direct investment in emerging markets—a question that sits at the intersection of two large literatures but had not been answered with credible cross-country causal evidence. Using an instrumental variables strategy that exploits the two-year-lagged log level of each country's commodity import price index as an instrument for reserve holdings, we show that the answer is yes.

The IV estimates imply an elasticity of approximately 1.85: a 10% increase in reserves raises inward FDI by about 18.5%. This is more than four times the fixed-effects OLS estimate, a gap that is consistent with downward bias arising from negative selection—countries tend to accumulate reserves precisely when unobserved risks are elevated and foreign investors retrench. Correcting for this bias reveals a causal effect that OLS systematically understates. The effect is not uniform across global conditions: the IV elasticity falls from 1.85 to 1.34 when crisis years (2008–2009 and 2020) are excluded, consistent with reserves functioning as insurance that is most valued during episodes of elevated global stress such as the global financial crisis and the COVID-19 shock, when downside risks are salient and external financing is most fragile.

The mechanism we emphasize is macro-financial credibility. Reserve buffers reduce the perceived probability of disruptive tail events—sharp devaluations, sudden imposition of capital controls, abrupt policy reversals—that are particularly costly for long-horizon, partially irreversible investments. By signaling external solvency and policy space, larger reserve stocks make emerging market destinations more attractive to multinational firms that must commit capital before the resolution of macro-financial uncertainty. The suggestive evidence in Section 6 is consistent with this channel: reserve accumulation is associated with lower sovereign risk premia and a reduced predicted probability of policy tightening, and conditioning on proxies for these risk dimensions attenuates the estimated reserve–FDI relationship in the expected direction.

These findings speak directly to a long-standing debate about the costs and benefits of reserve accumulation in emerging markets. The canonical cost estimate, due to Rodrik (2006), puts the social cost of holding reserves at roughly 1% of GDP annually, reflecting the spread between the return on reserve assets and the cost of external borrowing. This estimate has been influential in policy discussions about whether the large reserve buffers accumulated by emerging markets since the early 2000s represent an inefficient form of self-insurance. Our results suggest that this calculus is incomplete.

To the extent that reserves attract inward FDI—a relatively stable and productive source of external finance that also serves as a conduit for technology transfer and productivity spillovers—the benefits of reserve accumulation extend beyond crisis prevention to include effects on the longer-run investment environment. At the 2023 average values for our sample, the IV elasticity implies that a reserve increase from 22.45% to 24.7% of GDP would be associated with FDI inflows rising from 3.5% to approximately 4.18% of GDP in the near term—a gain of about 0.68 percentage points—and with cumulative FDI gains of roughly 1.75 percentage points of GDP over five years. Whether these investment gains are large enough to offset the cost estimate of Rodrik (2006) will depend on country-specific circumstances—the composition of FDI, the marginal productivity of foreign capital, and the opportunity cost of reserves—but the results make clear that any honest cost-benefit assessment of reserve policy must account for this channel.

The findings also have implications for how policymakers in emerging markets communicate reserve policy. If reserve buffers affect FDI partly through their effect on investor beliefs about macro-financial tail risk, then the signal sent by reserve accumulation may matter as much as the mechanical liquidity it provides. This suggests that transparency about reserve adequacy, and about the policy frameworks governing reserve use during stress episodes, may amplify the investment-attracting benefits identified here. Conversely, episodes of sharp reserve drawdown—even if warranted by fundamentals—may carry a signaling cost that is not fully captured by models focused on the mechanical insurance role of reserves.

Several limitations of our analysis point to productive directions for future work. Our sample covers 27 emerging market economies over 2001–2020, and the IV estimates should be interpreted as local average treatment effects: they capture the causal effect of reserve changes induced by external import-cost shocks, and the effect of purely discretionary reserve accumulation may differ. Future work with more granular data could explore heterogeneity across sectors—our mechanism predicts stronger effects for long-horizon, capital-intensive investments in manufacturing and infrastructure than for extractive FDI driven primarily by resource rents. Heterogeneity across exchange rate regimes and capital account arrangements is another natural dimension: reserves may be more effective signals of macro-financial stability in economies with more open capital accounts, where the threat of sudden stops is more salient. Finally, firm-level data on multinational investment decisions would allow a cleaner test of the mechanism by linking reserve buffers directly to entry and expansion decisions under uncertainty, rather than to aggregate net inflow data that conflates multiple margins of adjustment.

A Appendix

Table X: Variable definitions and sources

Variable	Definition and source
Dependent variable	
$\log(FDI/GDP)$	Natural logarithm of net FDI inflows as a percent of GDP (WDI), defined for $FDI/GDP > 0$.
Reserve measures	
$\log(res)$	Natural logarithm of total international reserves (IMF definition; WDI/IMF).
$\log(res/GDP)$	Natural logarithm of international reserves as a percent of GDP (constructed).
Controls	
$\log(GDPpc)$	Natural logarithm of GDP per capita, constant 2015 U.S. dollars (WDI).
$Growth$	Real GDP growth, annual percent (WDI).
$TradeOpen$	Trade openness: exports plus imports as a percent of GDP (WDI).
$Inflation$	Inflation, GDP deflator, annual percent (WDI).
$ExtBal$	External balance on goods and services as a percent of GDP (WDI).
$ExtDebt$	External debt stocks as a percent of GNI (WDI).
$\log(Spread)$	Natural logarithm of the sovereign spread over U.S. Treasuries (EMBI).
$Corruption$	Control of corruption index (Worldwide Governance Indicators).
$EconFreedom$	Economic Freedom Summary Index (Fraser Institute).
$KAOpen$	Capital account openness index (Chinn–Ito).
$CorpTax$	Statutory corporate income tax rate (OECD Corporate Tax Statistics).
$NatRes$	Total natural resource rents as a percent of GDP (WDI).
$Infrastructure$	Fixed telephone subscriptions per 100 people (WDI).
$HumanCapital$	Primary school enrollment, gross percent (WDI).
$LaborShare$	Labor share of income (Penn World Table).
$\log(ExpPriceIndex)$	Natural logarithm of the commodity export price index (IMF CTOT), used as a control.
Instrument	

Continued on next page

Table X – *Continued*

Variable	**Definition and source**
$\log(ImpPriceIndex)$	Natural logarithm of the commodity import price index (IMF CTOT; fixed historical import weights; deflated by manufactured-goods prices); used as an instrument.

B Sample composition and balanced-panel robustness

Our baseline analysis relies on an unbalanced country–year panel, where the estimation sample is pinned down by the joint availability of the outcome, the reserve measure, the import-price instrument, and the baseline control set. To make sample composition fully transparent, we report the exact country–year observations used by the baseline IV specification after listwise deletion (i.e., we drop any country–year with a missing value in the dependent variable, the lagged reserve regressor, the instrument, or any baseline control). Table XII lists the resulting country–year coverage, including the longest consecutive block for countries with at least 12 uninterrupted observations, while Table XIII reports the number of countries available in each year.

A potential concern with unbalanced panels is selection through sample entry/exit and intermittent missingness: if countries systematically drop in or out around major global shocks, or contribute only sporadically due to data gaps, changes in sample composition could mechanically affect the estimated relationship. To mitigate this concern, we re-estimate the baseline IV specification on a restricted subpanel that requires each included country to feature at least 12 consecutive years with no gaps in the baseline IV estimation sample. Table XII documents the countries that meet this requirement. Table XI reports the IV estimates on this consecutive-years subpanel. The reserve coefficients remain statistically significant and comparable in sign to the baseline, although somewhat attenuated in magnitude, indicating that the main results are not driven by changes in sample composition or by intermittent missingness in the unbalanced panel.

Table XI: IV Estimation: Reserves and Reserves-to-GDP (Consecutive-years subsample)

	(1) FS	(2) SS	(3) FS	(4) SS
$\log(M_i)_{t-2}$	1.498***		1.259***	
	(0.367)		(0.279)	
$\widehat{\log(Res)}_{t-1}$		1.634***		
		(0.409)		
$\log(\widehat{Res/GDP})_{t-1}$				1.944***
				(0.469)
Controls	Yes	Yes	Yes	Yes
Country and Year FE	Yes	Yes	Yes	Yes
Observations	227	227	227	227
First-stage F-stat (excluded instruments)	15.6		11.6	
Wu–Hausman p-value		0.176		0.141
Anderson–Rubin p-value ($H_0 : \beta = 0$)		0.0143		0.0143

Notes: Standard errors are two-way clustered by country and year. The reported first-stage F-statistic corresponds to the excluded instrument in the first stage. The Anderson–Rubin test is computed via the reduced-form regression and is robust to weak instruments.
**** p<0.01, ** p<0.05, * p<0.10.*

Table XII: IV estimation sample: country coverage and panel balance

	All countries		≥ 12 consecutive years	
Country	Years in sample	N	Longest consecutive block	Length
Argentina	2001–2020	20	2001–2020	20
Azerbaijan	2013–2020	8	—	—
Brazil	2013–2020	8	—	—
China	2002, 2007–2020	15	2007–2020	14
Colombia	2001–2003, 2005–2020	19	2005–2020	16
Costa Rica	2013–2020	8	—	—
Dominican Republic	2003–2020	18	2003–2020	18
Egypt, Arab Rep.	2002–2008, 2010, 2012–2015, 2017–2020	16	—	—
Gabon	2012, 2020	2	—	—
Georgia	2009–2020	12	2009–2020	12
India	2013–2020	8	—	—
Indonesia	2005–2020	16	2005–2020	16
Jamaica	2008–2020	13	2008–2020	13
Jordan	2012–2015, 2017–2020	8	—	—
Kazakhstan	2008–2020	13	2008–2020	13
Mexico	2001–2020	20	2001–2020	20
Mongolia	2013–2015, 2017–2020	7	—	—
Morocco	2001–2020	20	2001–2020	20
Peru	2001–2020	20	2001–2020	20
Philippines	2001–2010, 2015–2020	16	—	—
South Africa	2001–2020	20	2001–2020	20
Sri Lanka	2008–2010, 2016–2020	8	—	—
Thailand	2001–2003, 2005–2007	6	—	—
Tunisia	2003–2019	17	2003–2019	17
Turkiye	2001–2020	20	2001–2020	20
Ukraine	2001–2014, 2016–2020	19	2001–2014	14
Zambia	2013–2020	8	—	—

Notes: The table lists all 27 countries in the baseline IV estimation sample (365 country–year observations over 2001–2020). The right-hand columns report the longest consecutive block of years in the sample for countries with at least 12 uninterrupted observations.

Table XIII: Baseline IV estimation sample: number of countries by year.

Year	Countries in sample
2001	10
2002	12
2003	13
2004	11
2005	14
2006	14
2007	15
2008	17
2009	17
2010	18
2011	15
2012	18
2013	23
2014	23
2015	23
2016	22
2017	25
2018	25
2019	25
2020	25

C OLS robustness checks

Table XIV reports OLS robustness estimates using two reserve measures—reserves in levels and reserves scaled by GDP—and four transformations of FDI/GDP: log (baseline), IHS, shifted log, and levels, with winsorized variants for the latter two.

Table XIV: OLS robustness across reserve measures

Dep. var.: FDI/GDP	(1) Log	(2) Shifted log	(3) IHS	(4) IHS (wins.)	(5) Levels	(6) Levels (wins.)
Panel A: $\log(Res)_{t-1}$						
$\log(Res)_{t-1}$	0.419***	0.645**	0.592**	0.417***	2.804*	1.890**
	(0.124)	(0.241)	(0.211)	(0.124)	(1.393)	(0.807)
Observations	365	372	372	365	372	372
Adj. R^2	0.501	0.542	0.579	0.502	0.296	0.524
Panel B: $\log(Res/GDP)_{t-1}$						
$\log(Res/GDP)_{t-1}$	0.431***	0.580**	0.548***	0.429***	2.371**	1.755**
	(0.109)	(0.214)	(0.191)	(0.109)	(1.087)	(0.675)
Observations	365	372	372	365	372	372
Adj. R^2	0.502	0.536	0.575	0.504	0.286	0.521

Significance: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

D OLS placebo timing regressions

We estimate variants of the following specification:

$$y_{i,t} = \beta_0 \log(R_{i,t}) + \beta_{+1} \log(R_{i,t+1}) + \beta_{+2} \log(R_{i,t+2}) + \gamma' X_{i,t-1} + \mu_i + \tau_t + \varepsilon_{i,t}, \qquad (10)$$

where $y_{i,t} \equiv \log(FDI/GDP)_{i,t}$, μ_i and τ_t denote country and year fixed effects, and $X_{i,t-1}$ is the same lagged control set used in the baseline specifications. The reserve regressor is defined alternatively as $\log(\text{Res}_{i,t})$ (Panel A) or $\log(\text{Res}_{i,t}/GDP_{i,t})$ (Panel B).

Table XV reports the OLS placebo regressions. Columns (1)–(3) report single-horizon specifications including $\log(R_{i,t})$, $\log(R_{i,t+1})$, and $\log(R_{i,t+2})$ separately. Column (4) includes all three terms simultaneously.

Table XV: Placebo timing regressions: reserves and FDI (OLS)

Specification:	(1)	(2)	(3)	(4)
Panel A: Reserve levels, $\log(Res)$				
$\log(Res)_t$	0.695***			0.496***
	(0.150)			(0.152)
$\log(Res)_{t+1}$		0.624***		0.285
		(0.188)		(0.256)
$\log(Res)_{t+2}$			0.389**	-0.027
			(0.149)	(0.143)
Observations	365	365	365	365
Adj. R^2	0.532	0.523	0.500	0.532
Panel B: Reserves as % of GDP, $\log(Res/GDP)$				
$\log(Res/GDP)_t$	0.715***			0.693***
	(0.143)			(0.105)
$\log(Res/GDP)_{t+1}$		0.487***		0.291*
		(0.160)		(0.162)
$\log(Res/GDP)_{t+2}$			0.115	-0.419***
			(0.101)	(0.115)
Observations	365	365	365	365
Adj. R^2	0.535	0.508	0.485	0.540

*Significance levels: *** p<0.01, ** p<0.05, * p<0.10*

Columns (1)–(3) report single-horizon OLS placebo regressions for the contemporaneous (t), one-period-ahead ($t+1$), and two-period-ahead ($t+2$) reserve terms. Column (4) reports the joint OLS specification including all three terms simultaneously. Panel A uses reserve levels; Panel B uses reserves as a share of GDP.

Unlike the IV placebo results in Table V, the OLS estimates do not deliver a clean null on the reserve leads. This is expected and reflects the intrinsic limitations of OLS in this setting. Reserves are highly persistent: the within-country autocorrelation of $\log(R_{i,t})$ is close to unity, implying that $\log(R_{i,t-1})$, $\log(R_{i,t})$, $\log(R_{i,t+1})$ and $\log(R_{i,t+2})$ are nearly collinear. As a result, OLS cannot cleanly separate the predictive content of each horizon.

E Cumulative LP-IV responses

This appendix reports cumulative LP-IV impulse responses. We aggregate the horizon-specific elasticities $\widehat{\beta}_h$ from Section 5.4 into running sums that measure how the total effect of a reserve shock accumulates over time. For each horizon h, we define the cumulative effect as the partial sum of horizon-specific coefficients,

$$\text{CumEffect}(h) \equiv \sum_{j=1}^{h} \beta_j, \tag{11}$$

which measures how the reserve-driven effect on FDI builds up over successive years. In practice, we compute $\widehat{\text{CumEffect}}(h)$ by cumulating the estimated $\widehat{\beta}_j$ across horizons.

Figure III reports the cumulative responses for the same outcome transformations and reserve normalizations as the impulse response figure in the main text. Two patterns stand out. First, cumulative responses rise over the first few horizons, reflecting that impulse responses are concentrated in the short run. Second, the cumulative response remains positive and statistically significant through horizon $h = 5$ for the baseline log specifications, and through $h = 3$ for the alternative FDI transformations. The cumulative elasticity stabilizes at approximately 5 from $h = 5$ onward, implying a cumulative FDI/GDP gain of about 50% from a 10% reserve increase. This figure should be interpreted as an upper bound: the LP-IV estimates sum horizon-specific elasticities that are individually imprecise at longer horizons, as reflected in the widening confidence bands.

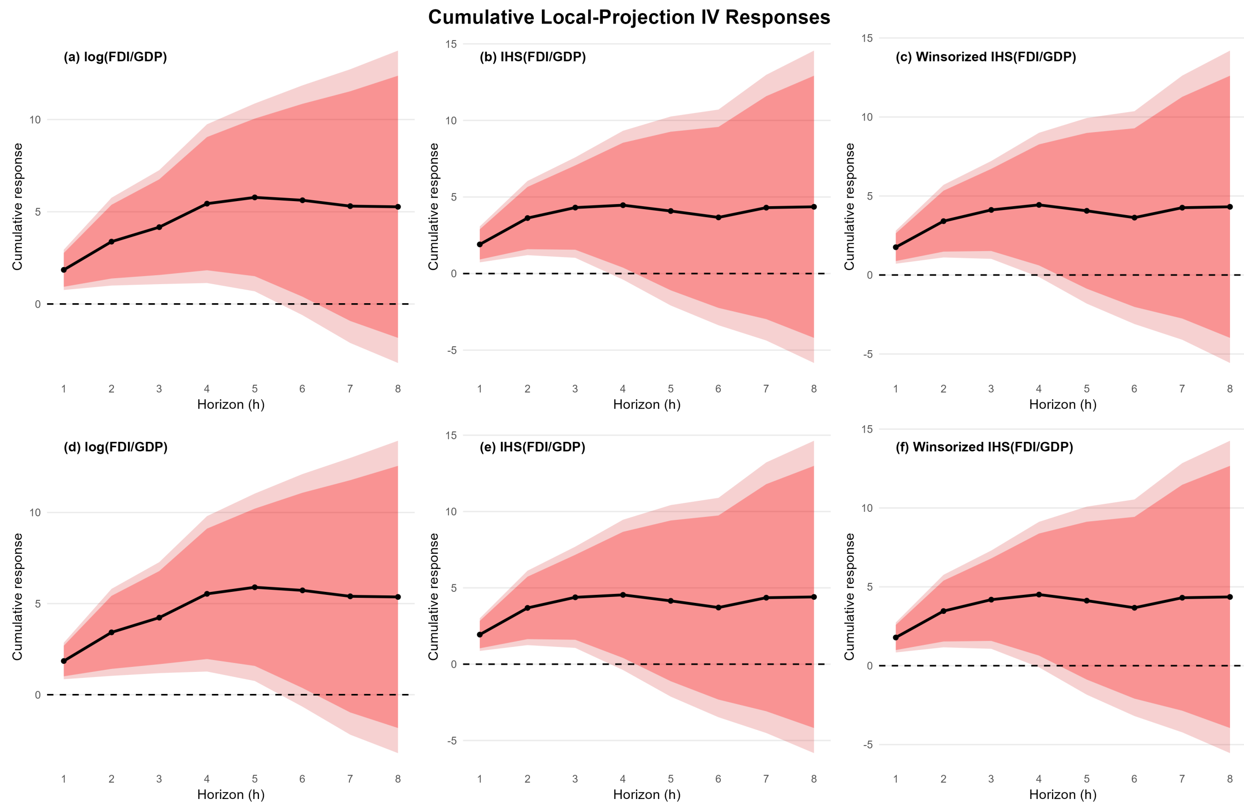


Figure III: **Cumulative LP-IV responses of inward FDI to reserves.** Notes: Cumulative LP-IV responses $\sum_{j=1}^{h} \widehat{\beta}_j$ $(h = 1, \ldots, 8)$ with 90/95% two-way clustered confidence bands.

F IV robustness checks and diagnostics

This appendix reports additional diagnostics and robustness checks for the instrumental-variables strategy. The goal is to document instrument relevance and to assess the stability of the IV estimates to alternative timing, sample restrictions, and inference procedures.

F.1 Alternative instrument lags

Our baseline instrument uses a two-year lag of the import commodity price index, $z_{i,t} = \log M_{i,t-2}$, to allow time for import-price shocks to affect balance-of-payments conditions and reserve management (one year) while limiting concerns about direct contemporaneous effects on FDI. Table XVI assesses the sensitivity of the 2SLS estimates to alternative timing choices by replacing $z_{i,t}$ with $\log M_{i,t-L}$ for $L \in \{1, 2, 3\}$, for both reserve measures.

Table XVI: Instrument timing: 2SLS estimates for $t-3$ to $t+2$

	$t-3$	$t-2$	$t-1$	t	$t+1$	$t+2$
Panel A: $\log(\widehat{Res})_{t-1}$						
$\hat{\beta}_{\text{IV}}$	1.884**	1.850***	4.718	-10.446	12.189	4.724
(SE)	0.863	0.557	4.175	20.076	67.151	11.100
First-stage F	16.4	19.8	3.2	0.2	0.0	0.2
Observations	365	365	365	365	365	365
Panel B: $\log(\widehat{Res/GDP})_{t-1}$						
$\hat{\beta}_{\text{IV}}$	1.799*	1.851***	5.513	-6.756	2.591	1.939
(SE)	0.878	0.510	4.174	5.683	3.551	3.898
First-stage F	17.8	19.5	2.3	0.6	0.4	1.4
Observations	365	365	365	365	365	365

Significance: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Notes: Each column reports a 2SLS estimate instrumenting the lagged reserve measure with the import-price index at the indicated timing $t+k$. The baseline specification uses $t-2$.

F.1.1 Targeted exclusions

We assess sensitivity to targeted sample exclusions. Large SDR allocations (2009) can mechanically affect measured reserves and potentially influence the first stage. We therefore re-estimate the baseline 2SLS excluding the SDR allocation year (2009), dropping all observations where the outcome year is 2009. Table XVII reports the corresponding 2SLS estimates and first-stage diagnostics for both reserve measures.

Table XVII: Targeted exclusions: sensitivity of 2SLS estimates

Sample restriction	$\hat{\beta}_{\text{IV}}$	(SE)	First-stage F	Obs.
Panel A: $\log(\widehat{Res})_{t-1}$				
Baseline (full IV sample)	1.850***	0.557	19.8	365
Exclude SDR allocation years (2009)	1.647***	0.507	14.7	315
Panel B: $\log(\widehat{Res/GDP})_{t-1}$				
Baseline (full IV sample)	1.851***	0.510	19.5	365
Exclude SDR allocation years (2009)	1.622***	0.329	14.9	315

Significance: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Notes: The exclusion drops all observations where the outcome year is 2009 (large SDR allocation).

F.1.2 Alternative clustering and leave-one-cluster-out (LOCO)

This subsection reports two compact diagnostics on the robustness of inference and influence. First, we recompute standard errors for the baseline IV specification under alternative clustering choices (two-way by country and year, country-only, and year-only). The point estimate is unchanged by construction, while statistical significance may vary with the clustering level. Second, we implement leave-one-cluster-out (LOCO) exercises, re-estimating the baseline IV specification while dropping one country (respectively, one year) at a time and reporting the range of the resulting IV coefficient.

Table XVIII shows results for both reserve measures. The associated standard errors imply significance at the 5% level under all clustering specifications in both cases. The LOCO ranges indicate that no single country or year drives the estimate: the coefficient remains positive and significant at 5% throughout for both definitions of reserves.

Table XVIII: Alternative clustering and leave-one-cluster-out (LOCO)

	$\widehat{\log(Res)}_{t-1}$		$\log(\widehat{Res/GDP})_{t-1}$	
	Estimate (SE)	Sig.	Estimate (SE)	Sig.
Alternative clustering				
Two-way clustering (country & year)	1.850 (0.557)	***	1.851 (0.510)	***
Country-only clustering	1.850 (0.638)	***	1.851 (0.637)	***
Year-only clustering	1.850 (0.610)	***	1.851 (0.559)	***
LOCO ranges for $\hat{\beta}_{IV}$				
Drop one country at a time	[1.428, 2.544]	sig. at 5% in all cases	[1.467, 2.678]	sig. at 5% in all cases
Drop one year at a time	[1.330, 2.945]	sig. at 5% in all cases	[1.318, 3.030]	sig. at 5% in all cases

Significance: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Notes: Alternative clustering rows report the baseline IV point estimate under different variance estimators. LOCO rows report the range of the IV coefficient across leave-one-cluster-out re-estimations; significance refers to the worst-case p-value across all exclusions.

F.2 Additional IV robustness: combining SDR allocations and import prices

As an additional robustness check, we strengthen the first-stage predictive content of the reserves instrumentation by augmenting the baseline IV set with an external policy-driven source of variation. Specifically, we instrument lagged reserves using (i) lagged IMF SDR allocation transfers scaled by constant GDP and transformed with the inverse hyperbolic sine, and (ii) lagged import-price conditions. We estimate a two-stage least squares specification with country and time fixed effects and two-way clustered standard errors (by country and year). We report overidentification diagnostics for the over-identified specification in the corresponding table; as usual under multi-way clustered inference, these tests should be interpreted with caution. The second-stage specification matches the baseline controls and lag structure.

This specification is reported solely as a robustness exercise; the preferred identification strategy and main results in the paper rely on the baseline instrument set. The purpose here is to verify that our conclusions are not sensitive to strengthening the first stage through an additional, plausibly exogenous source of variation in reserves. Reassuringly, the estimated IV elasticity of 1.6–1.7 is consistent with the baseline of 1.85, with the same sign and statistical significance. Overall, these results indicate that our main conclusions are robust to augmenting the instrument set with SDR-based variation and import-price shocks.

Table XIX: IV Estimation: Reserves and Reserves-to-GDP (Appendix: SDR + Import Prices Instruments)

	(1) FS	(2) SS	(3) FS	(4) SS
$(SDR_i)_{t-2}$	6.757***		10.773***	
	(0.641)		(2.202)	
$\log(M_i)_{t-2}$	1.435		1.433***	
	(0.866)		(0.425)	
$\widehat{\log(\text{Res})}_{t-1}$		1.706***		
		(0.531)		
$\log(\widehat{\text{Res/GDP}})_{t-1}$				1.598***
				(0.491)
Controls	Yes	Yes	Yes	Yes
Country and Year FE	Yes	Yes	Yes	Yes
Observations	365	365	365	365
First-stage F-test (excluded IVs)	10.00	10.00	10.10	10.10
Wu–Hausman p-value		0.0188		0.0307
Anderson–Rubin p-value ($H_0 : \beta = 0$)		6.81e-05		6.81e-05

*Notes: Columns (1)-(2) use reserves in levels; columns (3)-(4) use reserves-to-GDP. *** p<0.01, ** p<0.05, * p<0.10.*

G Alternative Definition of Policy Tightening

This appendix reports additional results obtained using an alternative definition of the policy tightening indicator. In contrast to the baseline specification, which imposes a threshold $\kappa = 0.99$, the alternative specification sets $\kappa = 1$, thereby classifying any non-positive change in the institutional index as a tightening episode.

Table XX: IV Estimation Results: Alternative Tightening Indicator ($\kappa = 1$)

Dep. var.: FDI/GDP	(1)	(2)
	$\widehat{\log(Res)}_{t-1}$	$\widehat{\log(Res/GDP)}_{t-1}$
$\widehat{\log(Res)}_{t-1}$	1.507**	
	(0.670)	
$\widehat{\log(Res/GDP)}_{t-1}$		1.475**
		(0.650)
$\widehat{\log(\text{Spread})}_{t-1}$	0.281	0.199
	(0.271)	(0.249)
$\widehat{Prob(\text{Str. Cap. Cont.})}_{t-1}$	0.095	-0.026
	(0.159)	(0.128)
Observations	331	331
Adj. R^2	0.448	0.442

Significance: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Notes: $\widehat{\log(\text{Spread})}_{i,t-1}$ and $\widehat{Prob(\text{Str. Cap. Cont.})}_{i,t-1}$ are fitted values from auxiliary regressions as described in Section 6. The alternative tightening indicator uses $\kappa = 1$ (any non-positive change in the institutional index).

H Robustness to Institutional Data Imputation

This section assesses the sensitivity of the baseline results to the imputation of missing values in the *Corruption* variable for 2001. We re-estimate the IV specification excluding all observations that rely on this interpolation, which also implies dropping 2002 due to lagged controls. As shown in Table XXI, coefficient magnitudes change only marginally, while statistical significance and the overall conclusions remain unaffected.

Table XXI: IV estimation excluding imputed 2001 values for Institutions I

	(1) FS	(2) SS	(3) FS	(4) SS
$\log(M_i)_{t-2}$	1.505***		1.486***	
	(0.391)		(0.441)	
$\widehat{\log(Res)}_{t-1}$		1.921***		
		(0.528)		
$\log(\widehat{Res/GDP})_{t-1}$				1.946***
				(0.508)
Controls	Yes	Yes	Yes	Yes
Country FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
Observations	353	353	353	353
First-stage F-stat (excluded instruments)	21.9		20.7	
Wu–Hausman p-value		0.00671		0.00717
Anderson–Rubin p-value ($H_0: \beta = 0$)		3.35e-05		3.35e-05

Notes: Columns (1)–(2) use reserves in levels; columns (3)–(4) use reserves-to-GDP. The sample excludes country-year observations in which the 2001 value of Institutions I was imputed using the average of 2000 and 2002. Standard errors are two-way clustered by country and year.

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Declaration of generative AI and AI-assisted technologies in the manuscript preparation process: During the preparation of this work the authors used Claude (Anthropic) and ChatGPT (OpenAI) in order to assist with code debugging and review, check spelling and grammar, improve clarity of exposition, and help with LaTeX formatting. After using these tools/services, the authors reviewed and edited the content as needed and take(s) full responsibility for the content of the published article.

This research did not receive any specific grant from funding agencies in the public, commercial, or not-for-profit sectors.

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