Trade, Multinational Production, and Regional Inequality

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Abstract

Globalization is often accompanied by declining barriers to both international trade and multinational production (MP), yet their joint implications for regional economies remain insufficiently understood. This paper examines how China's reductions in trade and MP barriers affected regional inequality in Korea, focusing on welfare and population distribution across regions. We first provide reduced-form evidence that regional exposure to MP, alongside trade, is an important and independent determinant of local employment. We then develop a multi-country, multi-sector general equilibrium model with trade, MP, and internal migration to quantify how China's cost reductions over 2000–2007 reshaped regional welfare and population allocation in Korea. The aggregate welfare gain is approximately 2.0%, with regional per-worker gains ranging from 1.7% in Seoul/Capital to 2.5% in the Southeast. Trade accounts for 1.4 percentage points of this gain, internal migration for 0.55 percentage points, and the vertical MP channel—through which affiliates abroad source inputs from the domestic parent—for approximately 0.04 percentage points. By jointly analyzing trade and MP, the paper shows that ignoring the MP channel would attenuate the estimated import-competition effect and misattribute part of the employment response to trade exposure alone.

Keywords: International trade, multinational production, regional inequality, China shock, internal migration

JEL Codes: F12, F14, F23, R12, R23

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1 Introduction

Over the past few decades, economic activity in Korea has become increasingly concentrated in the Seoul Metropolitan Area. Between 1990 and 2010, the area's shares of both population and gross regional product each rose by approximately ten percentage points. During the same period, China grew from a negligible partner[1] to Korea's largest counterpart in trade and multinational production (MP).[2] This paper asks whether China's rise contributed to Korea's increasing regional inequality, and whether multinational production—typically omitted from the analysis—is a key part of the answer.

A large literature examines how import competition from China reshapes regional labor markets (Autor et al., 2013; Topalova, 2010; Kovak, 2013). These studies treat trade as the sole channel through which China affects domestic outcomes. Yet China's integration into the global economy simultaneously reduced barriers to multinational production: as trade costs fell, firms also found it cheaper to relocate production stages abroad. The resulting outward MP operates through two opposing forces—it may substitute for domestic production, but it also generates demand for home activity through intra-firm sourcing of intermediate inputs from the parent. Crucially, trade and MP exposures are correlated across regions because both are driven by the same bilateral cost reductions. Analyzing import competition while omitting MP therefore creates omitted variable bias: the positive employment effect of vertical complementarity attenuates the negative import-competition coefficient, and may even reverse its sign. Controlling for MP is necessary to recover the causal effect of trade on regional labor markets.

Korea provides an ideal setting to study this joint channel. Between 2000 and 2007, outward MP to China expanded massively, with production by Korean affiliates reaching 15–25 percent of domestic output in sectors such as electronics, transport equipment, and machinery, while remaining below 5 percent in wood, paper, and basic metals.[3] Korean affiliates in China source roughly 20 percent of Korea's exports to China as intermediate inputs from their domestic parents,[4] creating a

[1] China's trade and inward FDI were heavily restricted prior to the early 1990s and were gradually liberalized in the run-up to WTO accession (Head and Ries, 1996; Whalley and Xin, 2010; Alviarez et al., 2022).

[2] We define MP as cross-border production activity, measured by the gross output of foreign affiliates. This differs from foreign direct investment (FDI), which measures capital flows; the two are related but distinct. A firm may sustain ongoing MP in a country long after the initial FDI transaction.

[3] Ratios computed from OECD AAMNE data, averaged over 2000–2007.

[4] Authors' calculations from OECD ICIO and Korean Customs data.

concrete, measurable vertical complementarity. At the same time, China's exports to Korea surged, generating large cross-regional variation in import competition. The coexistence of these forces—and their correlation through shared industrial composition—makes Korea a natural laboratory for identifying the MP channel alongside trade.

We proceed in two steps. First, we provide reduced-form evidence using a shift-share instrumental variable design that separately identifies the import-competition, export-opportunity, and outward-MP channels. This establishes the empirical relevance of MP as an independent determinant of regional employment. Second, we develop and quantify a multi-country, multi-sector general equilibrium model with trade, MP, and internal migration to evaluate how China's cost reductions—which we refer to as the China shock—reshaped regional welfare and population allocation.

Our model builds on Ramondo and Rodríguez-Clare (2013) and Galle et al. (2023), incorporating firm-level choices among domestic production, exporting, and MP, alongside the migration decisions of workers. The model links the China shock to regional inequality through three mechanisms. First, trade and MP cost reductions are asymmetric across sectors, generating industry-specific changes in labor demand. Second, workers sort across region-sector pairs in response: each worker's income depends on sectoral wages and an idiosyncratic productivity draw that varies across regions, reflecting, for example, region-specific vocational training or knowledge spillovers. Third, and most importantly for regional inequality, migration frictions impede this reallocation. Workers trapped in declining region-sector pairs cannot fully share in the gains accruing elsewhere, producing divergent welfare outcomes that depend on initial location. We calibrate the model to the year 2000, estimating trade and MP costs from OECD ICIO and AAMNE data and the vertical fragmentation parameter a^j following Alviarez (2019), and conduct two counterfactual exercises: comparing autarky to observed 2000 openness, and comparing the 2000 equilibrium to 2007 after the China-driven cost reductions.

Our main findings are as follows. In the reduced-form analysis, the preferred manufacturing stacked specification (KP $F = 8.81$) yields a positive and significant MP coefficient ($\hat{\beta}_{MP} = 0.034$, $p < 0.01$) alongside a negative import-competition coefficient ($\hat{\beta}_{IM} = -0.293$, $p < 0.05$). Omitting the MP control attenuates the import coefficient toward zero—and in the economy-wide crosssection, reverses its sign entirely—confirming the omitted variable bias we hypothesized.[5]

In the structural analysis, the 2000→2007 China shock raises aggregate Korean welfare by approximately 2.0%, with per-worker gains ranging from 1.7% in Seoul/Capital to 2.5% in the Southeast. A closed-form welfare decomposition (Proposition 1) separates this into trade (1.42 pp), internal migration (0.55 pp), and vertical MP (0.04 pp). Although the direct MP contribution is small, the MP channel matters substantially for accurate welfare measurement. A trade-only model understates the aggregate gain by 0.38 pp in the 2000→2007 exercise because it ignores vertical complementarity, but *overstates* by 0.13 pp in the autarky comparison because it ignores production displacement. This sign reversal reflects the net of the two opposing forces—complementarity and displacement—whose balance depends on the nature of the shock. Jointly analyzing trade and MP is therefore necessary both for causal identification and for accurate welfare accounting.

This paper contributes to three strands of the literature. First, we extend the literature on regional distributional effects of trade shocks (Topalova, 2010; Kovak, 2013; Autor et al., 2013; Acemoglu et al., 2016; Pierce and Schott, 2016; Dix-Carneiro and Kovak, 2017; Coşar and Fajgelbaum, 2016; Galle et al., 2023; Caliendo et al., 2019; Kleinman et al., 2023; Autor et al., 2021). Applied to Korea, Choi and Xu (2020) document manufacturing employment effects of Chinese import competition, Choi et al. (2024) show that the China shock reshaped internal migration, and Ahn and Choi (2020) provide evidence on vertical fragmentation through intermediate-input imports. We extend this work by incorporating the MP channel alongside trade, showing that omitting it attenuates the estimated import-competition effect, and by building a structural model that yields a welfare decomposition separating trade, MP, and migration gains.

Second, we contribute to the literature on home-country labor market consequences of outward MP (Helpman et al., 2004; Antràs and Helpman, 2004; Desai et al., 2009; Muendler and Becker, 2010; Harrison and McMillan, 2011; Ebenstein et al., 2014; Kovak et al., 2021; Boehm et al., 2020). A central econometric challenge is that MP and import competition are jointly determined, generating omitted variable bias when only one is controlled for. We show that this bias is empirically large for Korea and that the structural analog—the direction of welfare mismeasurement from omitting MP—depends on the nature of the counterfactual. The MP channel is not a nuisance to

[5] The economy-wide specification has low joint instrument strength (KP $F < 3$) and should be interpreted as a descriptive pattern rather than a causal estimate.

be controlled away, but a substantive force whose net effect shifts between complementarity and displacement as the shock varies.

Third, we extend the structural gains-from-trade-and-MP literature (Ramondo and Rodríguez-Clare, 2013; Irarrazabal et al., 2013; Tintelnot, 2017; Arkolakis et al., 2018; Head and Mayer, 2019) by adding regional heterogeneity through heterogeneous workers and internal migration (Artuç et al., 2010; Galle et al., 2023). Our Proposition 1 provides a closed-form welfare decomposition that nests the Arkolakis et al. (2012) sufficient statistic as a special case but adds a migration term $\hat{\mu}^k_{igg}$ with no counterpart in their representative-agent framework.

The rest of the paper is structured as follows. Section 2 presents the data sources and shift-share evidence. Section 3 develops the theoretical framework. Section 4 describes calibration, Section 5 presents counterfactual results, and Section 6 concludes.

2 Data and Empirical Strategy

This section details the data sources and empirical methodology used to quantify the joint impact of trade and multinational production (MP) on local labor markets in Korea. We first describe the data and document stylized facts on Korea's structural transformation alongside its deepening integration with China. We then develop a shift-share instrumental-variable design that separately identifies the import-competition, export-opportunity, and outward-MP channels, using supply-side shocks from other high-income countries as instruments. The economy-wide specifications serve as motivating evidence; our preferred specification for causal inference is the manufacturing-only stacked long-difference, which delivers robust identification of the vertical complementarity channel.

2.1 Data Sources and Processing

Our empirical analysis relies on three primary micro-datasets covering the period 2000 to 2020: establishment-level census data on local labor market outcomes, international production and trade data on supply-side shocks, and firm-level foreign investment data for granular heterogeneity analysis. A central challenge in analyzing long-term regional trends in Korea is the frequent rezoning of administrative boundaries. To ensure spatial consistency, we harmonize all geographic data to the 2000 administrative divisions, resulting in 230 consistent spatial units (Si/Gun/Gu) that serve as our definition of local labor markets.[6]

Local labor market outcomes are constructed using the Census on Establishments (2000–2020) from Statistics Korea. This census covers the universe of establishments in Korea with more than five employees and provides detailed annual information on employment, location, and industry classification. We map the native Korean Standard Industrial Classification (KSIC) codes to the ISIC Rev. 4 standard to ensure compatibility with international trade and investment data. Our primary outcome variable is the change in the manufacturing employment-to-population ratio, which captures the local industrial adjustment.

To measure multinational production (MP) and trade, we primarily utilize harmonized international data from the OECD. We obtain bilateral MP flows from the OECD Analytical AMNE (AAMNE) database, which provides the gross output of multinational affiliates by country of ownership and host location. This serves as our main measure of MP exposure (ΔMP). Bilateral trade flows are sourced from the OECD Inter-Country Input-Output (ICIO) Tables. These data allow us to construct symmetric, industry-level measures of exposure for both trade and investment between Korea, China, and a group of other high-income countries (O; see Section 2.3).

We complement these aggregate measures with a comprehensive dataset of outward foreign direct investment (FDI) transactions from the Export-Import Bank of Korea (Eximbank). This dataset records the date, destination, industry, and value of every outward investment made by Korean firms. We aggregate these transaction-level records to the industry-destination-year level to construct alternative MP proxies—such as "Net Investment" (gross investment minus repatriated capital) and the count of new foreign affiliates—which allow us to distinguish between the intensive and extensive margins of production relocation in our robustness checks. All monetary values are deflated to 2000 constant USD to account for inflation over the sample period.

[6]Major administrative changes during our sample period include the integration of Masan and Jinhae into Changwon City (2010), the merger of Cheongju City and Cheongwon County (2014), and the establishment of Sejong Special Self-Governing City (2012) from parts of Yeongi County and surrounding areas. We reconstruct consistent spatial units by aggregating all split or merged districts back to their year-2000 parent boundaries. For example, data for the current districts of Changwon (Uichang, Seongsan, Masanhappo, Masanhoewon, Jinhae) are aggregated into a single "Changwon" unit, and Sejong is re-mapped to the historical Yeongi and Cheongju boundaries to maintain a balanced panel of 230 regions.

2.2 Empirical Motivation: Korea's Integration with China and Structural Transformation

Korea's structural transformation over the past three decades exhibits three spatial patterns that motivate our empirical design and structural model.

First, aggregate deindustrialization coincided with—and was partly driven by—deepening integration with China. Figure 1 (Panel A) shows that as outward MP and exports to China expanded from the mid-1990s onward, the domestic manufacturing employment share steadily declined. Panel B reveals that this was not uniform contraction but compositional upgrading: traditional manufacturing employment stagnated while technology services exhibited explosive growth.

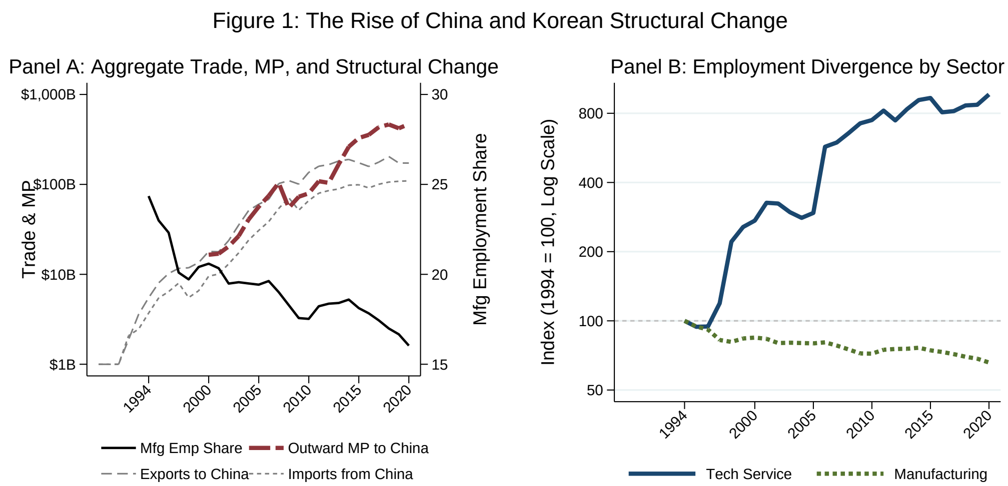

Figure 1: The Rise of China and Korean Structural Change

Notes: Panel A plots aggregate trade and multinational production against the manufacturing employment share. Panel B indexes employment in manufacturing and tech services to 1994=100 on a log scale.

Second, the aggregate decline masked dramatic spatial reallocation. Figure 2 maps the manufacturing employment share across municipalities in 1994 and 2020: regional hubs that once exceeded 50 percent manufacturing had largely faded by 2020. Meanwhile, knowledge-intensive activities grew into highly localized agglomerations. Figure 3 shows that ICT and R&D services—practically nonexistent in 1994—concentrated in a handful of metropolitan regions by 2020.

Third, the surviving domestic manufacturing base itself upgraded. Figure 4 shows that by 2020, high-tech manufacturing had clustered in the same regions where tech services agglomerated, whereas low-tech production—once widespread—had contracted sharply. This co-location of high-

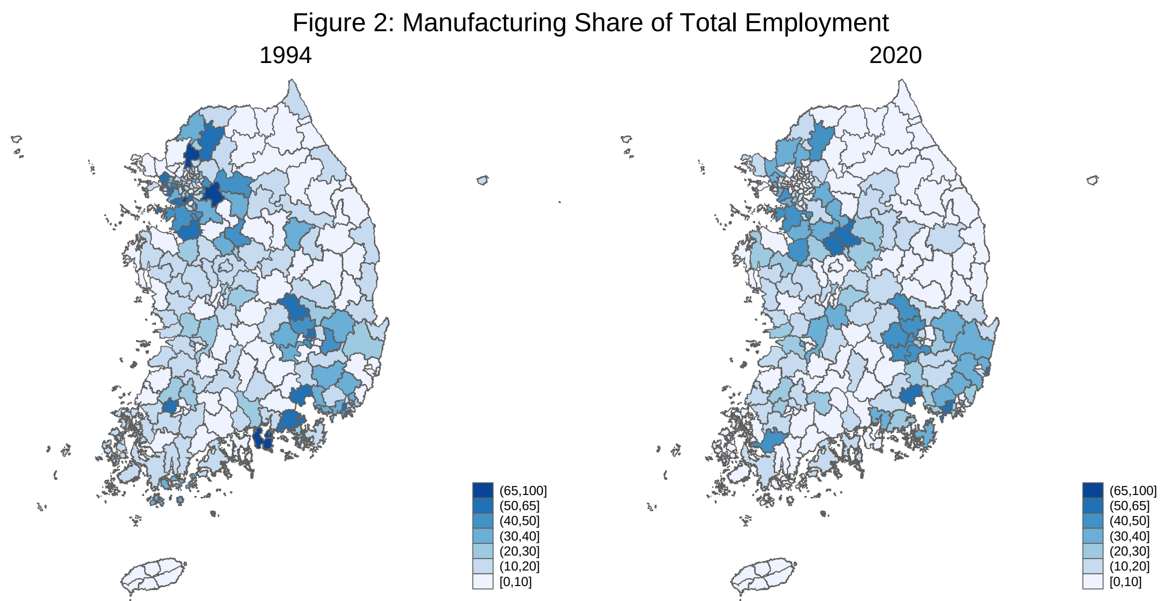

Figure 2: Manufacturing Share of Total Employment (1994 vs. 2020)

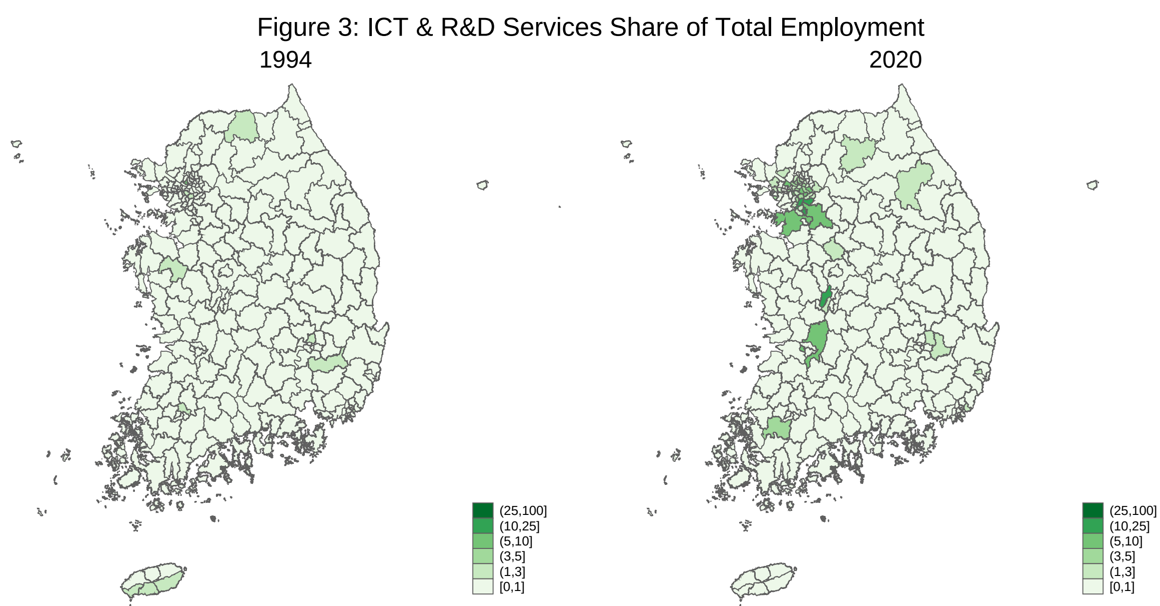

Figure 3: ICT & R&D Services Share of Total Employment (1994 vs. 2020)

tech manufacturing and upstream services is consistent with vertical complementarity: regions whose firms relocated labor-intensive stages to China retained headquarters functions and upstream intermediates, generating demand for skilled domestic activity rather than hollowing out.

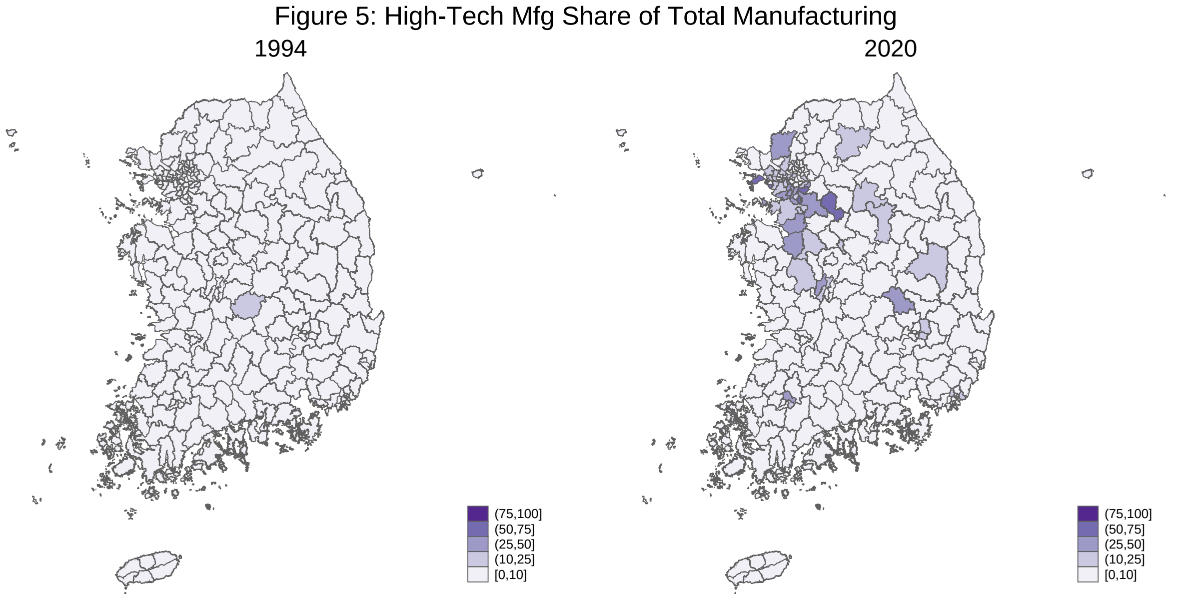

Figure 4: High-Tech Mfg Share of Total Manufacturing (1994 vs. 2020)

2.3 Empirical Strategy

To characterize the empirical relationship between trade and multinational production exposure and regional labor market outcomes, we employ a shift-share (Bartik) research design. Our strategy relies on the fact that regions differ in their initial industrial specialization, which exposes them differentially to aggregate shocks arising from China's integration into the global economy. This approach follows the framework of Autor et al. (2013) regarding trade exposure, while extending it to account for the variation in outward multinational production.

2.3.1 Econometric Specification

We estimate the effect of exposure to international markets on regional manufacturing employment using the following long-difference specification:

$$\Delta E_{r,t} = \alpha + \beta_{IM}\Delta IP_{r,t} + \beta_{EX}\Delta EP_{r,t} + \beta_{MP}\Delta MP_{r,t} + \gamma S_{r,t_0} + X'_{r,t}\delta + \epsilon_{r,t} \tag{1}$$

where $\Delta E_{r,t}$ represents the change in the manufacturing employment-to-population ratio in region r between the initial year t_0 and year t. The vectors $\Delta IP_{r,t}$, $\Delta EP_{r,t}$, and $\Delta MP_{r,t}$ denote the region's exposure to import competition, export opportunities, and outward multinational production, respectively. We do not include a corresponding term for inward multinational production (Chinese FDI into Korea) because Chinese inward MP in Korea was negligible during our sample period.[7]

We explicitly include the incomplete-shares control $S_{r,t_0} = \sum_j (L_{r,j,t_0}/L_{r,t_0})$—the region's initial manufacturing employment share—following Borusyak et al. (2025), which absorbs differences in the size of the exposed sector across regions and prevents the overall manufacturing concentration from confounding the shift-share estimates.

2.3.2 Measuring Regional Exposure

We construct regional exposure measures by interacting initial local industry shares with national-level changes in trade and investment flows. The shares $L_{r,j,t_0}/L_{r,t_0}$ use total regional employment in the denominator (economy-wide shares), which reflects the fact that trade shocks propagate to non-manufacturing regions through input-output and wage spillovers; restricting to manufacturing-only shares would understate exposure in regions with large upstream service employment. As a robustness check, we also construct shares using manufacturing employment only, providing a check that results are not driven by the inclusion of non-tradable industries.[8] The exposure to Chinese import competition, $\Delta IP_{r,t}$, is defined as:

$$\Delta IP_{r,t} = \sum_j \frac{L_{r,j,t_0}}{L_{r,t_0}} \left(\frac{\Delta M_{j,t}^{CHN \to KOR}}{L_{j,t_0}} \right) \tag{2}$$

where the term in parentheses represents the national-level shock: the change in Korean gross imports from China in industry j, normalized by base-period national employment *in sector* j, L_{j,t_0}.[9] $\Delta M_{j,t}^{CHN \to KOR}$ is the gross bilateral trade flow from China to Korea in sector j, sourced from OECD ICIO tables. Since bilateral intra-firm import flows are not separately observable in these data, gross trade is used throughout; this is standard in the shift-share trade literature (Autor

[7] Consequently, we cannot construct a meaningful regional variation for inward MP exposure. OECD AAMNE data shows that Chinese affiliates' share of production in Korea remained near zero throughout the 2000–2020 period.

[8] For the main long-difference specification (2000–2020), the base-year shares $L_{r,j,t_0}/L_{r,t_0}$ are measured in 2000 ($t_0 = 2000$). For the stacked first-difference robustness (Section 2.3.1), each 10-year period uses its own lagged base year: 2000 shares for the 2000–2010 period and 2010 shares for the 2010–2020 period. Using period-specific shares in the stacked model controls for changes in industrial composition between decades.

[9] L_{j,t_0} is national employment in sector j at the base year, following Autor et al. (2013); dividing by sector-level (not aggregate) employment ensures the national shock measures the change in imports per initial sector-worker and does not conflate industry-specific exposure with economy size.

et al., 2013). Symmetrically, we define exposure to Outward MP, $\Delta MP_{r,t}$, using the change in the sales of Korean affiliates in China ($\Delta MP_{j,t}^{KOR \rightarrow CHN}$), sourced from the OECD AAMNE database:

$$\Delta MP_{r,t} = \sum_j \frac{L_{r,j,t_0}}{L_{r,t_0}} \left(\frac{\Delta MP_{j,t}^{KOR \rightarrow CHN}}{L_{j,t_0}} \right) \tag{3}$$

This measure captures the intensive margin of production relocation. A region initially specialized in industries where Korean affiliates subsequently expanded production in China will exhibit a high value of $\Delta MP_{r,t}$. Symmetrically, exposure to export opportunities, $\Delta EP_{r,t}$, is defined as:

$$\Delta EP_{r,t} = \sum_j \frac{L_{r,j,t_0}}{L_{r,t_0}} \left(\frac{\Delta X_{j,t}^{KOR \rightarrow CHN}}{L_{j,t_0}} \right) \tag{4}$$

where $\Delta X_{j,t}^{KOR \rightarrow CHN}$ is the change in Korean exports to China in industry j, normalized by base-period national employment.

2.3.3 Identification: Relevance and Exclusion

Estimating Equation (1) via OLS may yield biased estimates if domestic Korean demand shocks drive both the outcome and the trade flows. To address this, we identify supply-side shocks originating in China.

Relevance. Economically, China-side supply forces (productivity growth, trade-cost declines, policy liberalization) that raise exports to Korea or attract Korean outward MP also raise the corresponding flows with other high-income countries. Following Autor et al. (2013), we construct the instrument from a group of other high-income countries (O) exposed to the same China-side supply shock through their own bilateral channels. Specifically, O comprises six economies—the United States, Japan, Germany, Taiwan, the Netherlands, and Singapore—selected as the intersection of (i) countries with bilateral MP data available in the OECD AAMNE and (ii) major global value chain (GVC) partners with substantial trade and investment linkages to China. These countries collectively account for approximately 40% of global merchandise trade and are the primary host destinations for Korean outward MP outside China. Developing economies are excluded because they may compete directly with China in third markets, violating the relevance condition.

The same six-country basket is used for all three instrument channels (import, export, and MP); as a robustness check, we re-estimate the baseline after dropping the United States, Japan, or both—the countries whose business cycles are most synchronized with Korea's—from the basket (Appendix A.10).

We instrument the Korean shocks ($\Delta M^{CHN \to KOR}$ and $\Delta MP^{KOR \to CHN}$) with their other-high-income-country analogs ($\Delta M^{CHN \to O}$ and $\Delta MP^{O \to CHN}$). We document the relevance of these instruments using the Sanderson and Windmeijer (2016) partial F-statistic for each endogenous regressor and the Kleibergen and Paap (2006) rk F-statistic (KP F) for joint relevance.

Exclusion and Threats to Identification. Validity of the shift-share design requires that the industry-level instruments be orthogonal to the structural error term in the outcome equation (Borusyak et al., 2022):

$$\mathbb{E}\big[g_j^{\mathrm{IV}}\,\bar{\varepsilon}_j\big] = 0 \quad \text{where} \quad \bar{\varepsilon}_j = \sum_r s_{rj}\varepsilon_r$$

Threats to exclusion arise when forces other than the China-side supply move with both Korean regional outcomes and the O-country instruments. A primary concern is global demand shocks—for example, a global technology boom that raises demand for electronics in both Korea and the O countries, thereby simultaneously increasing Chinese exports to both.

We address these threats in three ways. First, we attenuate demand linkages by verifying that the MP coefficient is stable when dropping the United States, Japan, or both from the O basket, and when replacing them with diversified alternatives whose multinational production in China is not concentrated in automotive industries (Appendix A.10). Second, we condition on the incomplete-shares control S_r to prevent the results from being driven by the secular decline of manufacturing. Third, unlike prior studies that treat MP as part of the error term, we explicitly model outward MP as an endogenous exposure and instrument it with O-country analogs and alternative proxies (investment counts and firm numbers). All baseline specifications are exactly identified (one instrument per endogenous variable), so no overidentification test is available; the robustness of results across alternative instrument choices provides indirect evidence of instrument validity.

Pre-trend test. A key assumption for shift-share identification is that the 2000 industry shares are pre-determined with respect to the 2000–2020 outcome. We test this by regressing pre-period regional employment growth (1994–2000) on the same shift-share exposures used in the main specification.[10] Under the null of no pre-trend, the coefficients should be statistically indistinguishable from zero.

We run two distinct pre-trend tests. The first (Table 17, Appendix A.1) regresses 1994–2000 regional employment growth on the endogenous *exposure* variables themselves. In the joint specification, all three exposure coefficients are statistically insignificant ($\hat{\beta}_{IM} = -0.255$, $\hat{\beta}_{EX} = 0.602$, $\hat{\beta}_{MP} = 0.057$; none significant at 10%), indicating that regional industry composition in 2000 does not predict pre-period employment patterns once all three channels are conditioned on jointly.

The second test (Table 16, Appendix A.1) checks whether the O-country *instruments* themselves predict 1994–2000 employment growth—a direct test of the exclusion restriction. The MP and export instruments are clean: the MP instrument becomes insignificant in the joint specification ($p = 0.196$), and the export instrument is insignificant throughout. The import instrument, however, remains significant at 5% in the joint specification ($\hat{\beta} = 0.185^{**}$), indicating a pre-existing correlation between import-competition-exposed regions and declining pre-period employment. This pattern is consistent with the broader China-shock literature (Autor et al., 2013) but represents a genuine limitation for causal inference on β_{IM} under the O-country analog.

We address this directly. China's pre-committed WTO accession tariff reductions provide an alternative instrument for the import channel that is, by construction, pre-determined: the tariff schedule was legally fixed before our sample period and reflects supply-side commitments independent of Korean regional conditions. Under this alternative IV (Appendix A.9), the import displacement result is confirmed and the MP complementarity estimate is unchanged ($\hat{\beta}_{MP} = 0.081^{**}$). The Hansen J over-identification test ($p = 0.810$) confirms that the WTO tariff IV and the O-country import analog are jointly consistent, implying the pre-trend in the O-country instrument does not materially bias the structural coefficients. Causal claims on β_{IM} in the main tables should be read with the caveat that the O-country import instrument has a pre-trend; claims on β_{MP} are unaffected.

[10] We use 1994, the earliest available year in the Establishment Census, as the pre-period start. The pre-period outcome is $\Delta E_{r,1994\to2000}/E_{r,1994}$.

Stacked First-Differences. Finally, a single long difference (2000–2020) could confound import exposure with decade-specific shocks or slow regional trends. To exploit within-region changes across subperiods while absorbing aggregate shocks, we estimate a stacked first-difference model (Autor et al., 2013):

$$\Delta E_{r,\tau} = \sum_c \beta_c \Delta X_{r,\tau}^{(c)} + \gamma S_{r,\tau} + \delta_r + \lambda_\tau + \varepsilon_{r,\tau} \tag{5}$$

where δ_r represents region fixed effects and λ_τ represents period fixed effects. This specification isolates variation in the *acceleration* of trade and MP exposure within regions over time. Standard errors are clustered at the region level to account for serial and spatial correlation.

Table 1 reports summary statistics for the key variables.

Table 1: Summary Statistics

	Long-Difference Sample (2000–2020)				
	Mean	SD	Min	Max	N
Dependent variable					
ΔE_r (10-year employment growth rate)	0.571	0.317	−0.023	1.695	230
Endogenous exposure variables (per worker, \$1,000s)					
$\Delta IP_r^{\text{CHN}\to\text{KOR}}$ (Import competition)	0.687	0.579	−0.899	4.817	230
$\Delta EP_r^{\text{KOR}\to\text{CHN}}$ (Export exposure)	0.491	0.450	0.089	3.786	230
$\Delta MP_r^{\text{KOR}\to\text{CHN}}$ (Outward MP)	4.438	4.258	0.687	20.471	230

Notes: Exposure variables are Bartik-weighted sums of industry-level shocks, normalized by initial regional employment. The dependent variable is the symmetric growth rate of regional employment over the indicated time horizon. The sample comprises 230 Si/Gun/Gu administrative units matched across the 2000 and 2020 Establishment Census.

2.4 Empirical Results

Table 2 presents the estimation results for the long-difference specification (2000–2020) outlined in Equation (1). Columns (1)–(2) report economy-wide exposure results and Columns (3)–(4) report manufacturing-only exposure results. Within each panel, the first column includes trade channels only and the second adds outward MP, yielding the preferred full specification. All columns report two-stage least squares (2SLS) estimates with domestic shocks instrumented using O-country analogs, and include province fixed effects and the incomplete-shares control S_{r,t_0}. We present the economy-wide specifications as motivating patterns; given the low joint Kleibergen–Paap F-statistics (2.35 and 1.66), formal causal inference in the long-difference rests on the individual Sanderson–Windmeijer partial F-statistics and the Anderson–Rubin confidence sets (Appendix A.5). Our preferred specification for causal claims is the manufacturing-only stacked long-difference (Table 3, Column 2), which achieves joint $F = 8.81$.

2.4.1 Main Results: Trade and Vertical Complementarity

We begin by examining the impact of trade exposure in isolation. In the economy-wide specification (Column 1), the import competition coefficient is positive but statistically imprecise ($\beta_{IM} = 0.325$, s.e. 1.124). In the manufacturing-only specification (Column 3), the import coefficient turns negative ($\beta_{IM} = -0.081$), consistent with trade-induced displacement in the exposed tradable sector, though it remains imprecise. The trade-only specifications provide limited explanatory power on their own.

Table 2: Main Results: Shift-Share IV Estimation (2000–2020)

	Economy-wide Exposure		Manufacturing Exposure	
	(1) Trade	(2) Full	(3) Trade	(4) Full
$\beta^{\text{IM}}_{\text{CHN}\to\text{KOR}}$	0.325	-0.933		
	(1.124)	(1.186)		
$\beta^{\text{EX}}_{\text{KOR}\to\text{CHN}}$	-0.142	0.848		
	(1.409)	(1.415)		
$\beta^{\text{MP-out}}_{\text{KOR}\to\text{CHN}}$		0.093***		
		(0.030)		
$\beta^{\text{IM}}_{\text{CHN}\to\text{KOR}}$ (Mfg)			-0.081	0.101
			(0.407)	(0.463)
$\beta^{\text{EX}}_{\text{KOR}\to\text{CHN}}$ (Mfg)			0.176	-0.217
			(0.494)	(0.560)
$\beta^{\text{MP-out}}_{\text{KOR}\to\text{CHN}}$ (Mfg)				0.031***
				(0.010)
Observations	230	230	230	230
Number of Shocks				
Kleibergen-Paap F	6.83	2.35	2.20	1.66
SW F: Import		12.01		5.05
SW F: Export		13.78		4.99
SW F: MP-out		17.80		93.80

Standard errors are robust. All regressions include region fixed effects.
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Crucially, adding outward MP substantially changes the picture. In the full economy-wide specification (Column 2), the outward MP coefficient is positive and significant ($\beta_{MP} = 0.093^{***}$, s.e. 0.030), while the trade coefficients remain imprecise. The full manufacturing specification (Column 4) delivers a similar pattern ($\beta_{MP} = 0.031^{***}$, s.e. 0.010). These estimates challenge the popular "hollowing out" narrative and are consistent with *vertical complementarity*: as Korean firms expand production abroad, they increase demand for headquarters services and upstream intermediates produced at home. The causal interpretation is most firmly established in the manufacturing stacked specification (Table 3, Column 2, KP $F = 8.81$), which yields $\beta_{MP} = 0.034^{***}$.

Instrument diagnostics. The joint Kleibergen–Paap F-statistics are 2.35 (economy-wide) and 1.66 (manufacturing), below the conventional threshold of 10. The economy-wide long-difference estimates should therefore be read as descriptive patterns. However, the Sanderson–Windmeijer (SW) partial F-statistics—which test individual instrument relevance conditional on the other channels (Sanderson and Windmeijer, 2016)—confirm that each instrument is individually strong. In the economy-wide specification, SW $F = 12.01$ (import), 13.78 (export), and 17.80 (MP-out), all above 10. In the manufacturing specification, SW $F = 5.05$ (import), 4.99 (export), and 93.80 (MP-out), with the MP instrument particularly strong. Anderson–Rubin confidence sets, which provide inference robust to weak instruments, are reported in Appendix A.5; the AR set for β_{MP} in the manufacturing specification is $[0.007, 0.054]$, excluding zero.

2.4.2 Stacked First-Differences

To ensure that our results are not driven by the specific endpoints of the long-difference (2000 and 2020) or by slow-moving regional trends, we estimate a stacked first-difference model over two ten-year periods (2000–2010, 2010–2020). This specification includes region- and period-fixed effects, identifying parameters solely from within-region variation in the acceleration of shock exposure.

Table 3 confirms the robustness of our baseline findings. The economy-wide stacked specification (Column 1) reports KP $F = 0.79$, indicating severe instrument weakness; we focus on the manufacturing-only results (Column 2). In this preferred specification, the import competition coefficient is negative and significant ($\beta_{IM} = -0.293^{**}$), confirming trade-induced displacement, the export coefficient is positive ($\beta_{EX} = 0.113^{***}$), and the outward MP coefficient remains positive and significant ($\beta_{MP} = 0.034^{***}$).

Table 3: Stacked First-Difference Results (10-Year Windows)

	Stacked Panel (10-Year)	
	(1) Economy-wide	(2) Mfg-Only
$\beta^{\text{IM}}_{\text{CHN}\to\text{KOR}}$	-1.752	
	(1.374)	
$\beta^{\text{EX}}_{\text{KOR}\to\text{CHN}}$	0.748	
	(0.527)	
$\beta^{\text{MP-out}}_{\text{KOR}\to\text{CHN}}$	0.145	
	(0.109)	
$\beta^{\text{IM}}_{\text{CHN}\to\text{KOR}}$ (Mfg)		-0.293**
		(0.119)
$\beta^{\text{EX}}_{\text{KOR}\to\text{CHN}}$ (Mfg)		0.113***
		(0.040)
$\beta^{\text{MP-out}}_{\text{KOR}\to\text{CHN}}$ (Mfg)		0.034***
		(0.013)
Observations	465	465
Kleibergen-Paap F	0.79	8.81
SW F: Import	2.91	30.58
SW F: Export	2.95	32.47
SW F: MP-out	2.75	28.63

Standard errors clustered by region. Region and Year FE included.

KP F: joint Kleibergen-Paap rank statistic (3 endog). SW F: Sanderson-Windmeijer partial F for each channel.

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

Notes: Column (1) (economy-wide) reports KP $F = 0.79$; instrument weakness is severe and Column (1) is not used for causal inference. Column (2) (manufacturing-only) reports joint KP $F = 8.81$, below the Stock–Yogo critical value for three endogenous variables (≈ 13.91 at 10% size distortion). The Sanderson–Windmeijer partial F-statistics—reported in the lower rows of the table—confirm that each channel is individually well-identified conditional on the others: SW $F = 30.6$ (import), 32.5 (export), 28.6 (MP-out), all well above 10. The low joint KP F reflects correlation among the three O-country instruments, not individual instrument weakness.

The joint KP $F = 8.81$ falls below the Stock–Yogo critical value for three endogenous variables (≈ 13.91 at 10% size distortion). As with the long-difference estimates, however, the SW partial F-statistics confirm strong individual instrument relevance: SW $F = 30.6$ (import), 32.5 (export), and 28.6 (MP-out), all well above 10. The low joint KP F reflects correlation among the three O-country instruments, not individual weakness. The Anderson–Rubin confidence set for β_{MP} is [0.022, 0.060], excluding zero (Appendix A.5); the AR sets for the trade channels are uninformative, reflecting their comparatively lower individual F-statistics. The MP complementarity result is thus robust across both the long-difference and stacked designs.

Economic magnitude. To interpret these coefficients in economically meaningful units, we compute standardized effects using the cross-regional standard deviations of the exposure variables (Table 1).[11] In the preferred manufacturing stacked specification, a one-standard-deviation increase in import competition exposure (SD = 0.58) is associated with a 17.0 percentage point *decline* in ten-year regional employment growth—roughly 30% of the cross-regional mean. Conversely, a one-standard-deviation increase in outward MP exposure (SD = 4.26) is associated with a 14.5 percentage point *increase* in employment growth, or approximately 25% of the mean. The export channel is smaller: a one-standard-deviation increase (SD = 0.45) translates to a 5.1 percentage point gain, about 9% of the mean. These magnitudes confirm that outward MP is not merely statistically significant but economically large: its standardized effect ($\tilde{\beta}_{MP} = 0.46$) is comparable in absolute value to the import-competition channel ($\tilde{\beta}_{IM} = -0.54$), underscoring the quantitative importance of omitting MP from standard trade-shock analyses.

Robustness and validity. We conduct six validity checks on the shift-share design, with full results in Appendix A.

(i) Shock-level inference. Following Borusyak et al. (2025), we re-estimate at the industry level to confirm that findings are driven by variation in exposure shocks rather than spurious regional trends (Appendix A.2).

(ii) Rotemberg weights and C29 concentration. Identification is heavily concentrated in Motor Vehicles and Trailers (C29), which accounts for 86% of the MP-out Rotemberg weight ($N_{\text{eff}} = 1.2$),

[11] The standard deviations are computed from the long-difference sample ($N = 230$). Stacked-panel standard deviations are similar due to the within-region differencing.

reflecting Korea's large-scale automotive supply-chain integration with China (Table 19). This concentration warrants scrutiny: the complementarity finding could reflect the automotive sector's inherently high upstream content rather than a general mechanism. Dropping all C29 industries from the exposure measure (Table 20, Appendix A.3), the MP coefficient remains positive ($\hat{\beta}_{MP} = 0.035$) but loses significance (s.e. = 0.258) as the KP F-statistic drops to 0.10—reflecting the loss of the dominant source of identification rather than a sign reversal. The bias-weighted F-statistic in the baseline is 37.0, confirming that the automotive-sector variation, though concentrated, is strongly relevant. To assess whether the concentration is an artifact of the auto-heavy United States and Japan in the O basket, we construct alternative baskets replacing them with France, Switzerland, and Austria—countries whose Chinese MP is concentrated in chemicals and pharmaceuticals rather than motor vehicles. The MP coefficient remains positive and significant across all seven compositions ($\hat{\beta}_{MP} \in [0.022, 0.055]$), though the C29 Rotemberg weight declines only from 86% to 72% ($N_{\text{eff}} = 1.8$), confirming that the concentration is partly structural—rooted in the automotive sector's outsized role in Chinese outward MP—rather than driven by baseline country selection (Appendix A.10).

(iii) First-stage relationships. Industry-level first-stage regressions confirm strong relationships between Korean shocks and the O-country instruments ($t > 3$), validating the relevance of our supply-side strategy (Table 22).

(iv) Manufacturing-only shares. Constructing exposure shares solely from manufacturing employment yields a slightly larger positive MP coefficient ($\beta \approx 0.13$), confirming that results are not driven by the inclusion of non-tradable industries.

(v) WTO tariff instrument. A concern is that the O-country import analog may partly reflect Chinese demand cycles rather than supply-side forces. We construct an alternative instrument based on China's pre-committed WTO accession tariff reductions (December 2001), which were negotiated before our sample period and are exogenous to any individual trading partner's demand conditions. Under this alternative IV, the MP coefficient ($\hat{\beta}_{MP} = 0.081^{**}$, s.e. 0.034) is stable. In the over-identified specification using both the WTO tariff Bartik and the O-country import analog, the Hansen J p-value is 0.810, failing to reject joint validity (Appendix A.9).

(vi) Alternative MP proxies. Using Eximbank transaction-level data, we find a striking heterogeneity across MP margin definitions (Table 24). Count-based extensive-margin measures (number

of new affiliates) are positively and significantly associated with domestic employment, whereas value-based intensive-margin measures (gross and net investment) yield near-zero and imprecisely estimated coefficients. This suggests that affiliate entry—rather than the scale of capital flows—is the margin most strongly linked to domestic employment complementarity.

Cross-regional spillovers (SUTVA). The shift-share design interprets $\hat{\beta}$ as the causal effect of exposure on the treated region's employment, which requires that one region's treatment does not affect other regions' outcomes (SUTVA). This assumption is violated in our setting through at least two channels. First, internal migration redistributes workers across regions in response to local shocks, so a region's employment outcome depends partly on the shocks experienced by other regions (Caliendo et al., 2019). Second, input-output linkages between regions transmit sector-specific shocks along supply chains, so an upstream region's import-competition shock may reduce intermediate demand from downstream regions. Both channels imply that the shift-share coefficients capture *partial-equilibrium* effects that hold other regions' exposures fixed, rather than the total effect of a nationwide shock. The structural model in Section 3 addresses this directly: it solves for the full general-equilibrium allocation in which all regions and sectors interact simultaneously, so cross-regional spillovers through migration and trade are endogenous. The reduced-form estimates should therefore be interpreted as identifying the direction and relative magnitude of each channel, while the GE model provides the welfare accounting that incorporates spillovers.

2.5 From Reduced-Form Evidence to Structural Analysis

The shift-share estimates in Section 2.4 establish three motivating regularities that guide the structural model developed below.

(i) Multinational production is a distinct and quantitatively important channel. Outward MP exposure is positively and robustly associated with regional employment growth, independent of trade exposure. This pattern holds across both the long-difference and stacked specifications, across economy-wide and manufacturing-only exposure measures, and across two independent data sources for trade and investment. The causal interpretation is most firmly supported by the manufacturing stacked specification, where the Anderson–Rubin confidence set for the MP coefficient excludes zero. A model that accounts only for trade would be mis-specified: it would attribute to trade effects that are partly—and, in our estimates, primarily—driven by the cross-border reorganization of production by multinational firms.

(ii) A trade-only model mismeasures the welfare effects of China's integration. The positive employment effect of outward MP implies that as Korean firms expand abroad, they pull resources through supply chains rather than releasing them. A trade-only welfare model ignores this backward linkage and overstates how much domestic production is displaced by imports; at the same time, it ignores that affiliates abroad replace some domestically produced output, dissipating gains from trade liberalization. The net direction of mismeasurement depends on the nature of the shock—whether it primarily reduces trade costs or MP costs—and cannot be signed from reduced-form estimates alone. A structural model that endogenizes both channels is therefore necessary for welfare accounting; Section 5 quantifies the magnitude of this mismeasurement.

(iii) The relevant margin of MP is organizational, not financial. The Eximbank evidence reveals that extensive-margin affiliate entry—the establishment of new foreign subsidiaries—drives the domestic employment complementarity, whereas intensive-margin capital flows are uninformative. This heterogeneity motivates a model in which firms allocate production across locations according to comparative advantage, with multinational production governed by entry costs rather than capital flows. The structural model should therefore feature firm-level location choice with vertical fragmentation—an affiliate that sources intermediates from its parent—rather than treating MP as a simple capital reallocation.

Together, these three facts motivate a model that is multi-country, multi-sector, and explicitly incorporates both trade and MP as endogenous margins—the framework to which we now turn.

3 The Model

We develop a multi-country, multi-sector general equilibrium model with international trade, multinational production (MP), and internal migration. The model combines the trade and MP framework of Ramondo and Rodríguez-Clare (2013) with the internal migration framework of Galle et al. (2023), adding multi-sector input-output linkages as in Caliendo and Parro (2015). Relative to each building block, we make three extensions. First, whereas Ramondo and Rodríguez-Clare (2013) features a single sector and no worker heterogeneity, we introduce J sectors with Cobb-Douglas

IO linkages and heterogeneous workers who sort across region-sector pairs. Second, whereas Galle et al. (2023) models migration in response to trade shocks only, we add the MP margin so that firms' location choices feed back into local labor demand and hence migration incentives. Third, whereas Caliendo and Parro (2015) models trade with IO linkages but no MP, we embed the full $\{i, l, n\}$ origin-production-destination structure so that multinational affiliates source inputs from both home and host countries. The resulting model yields Proposition 1, a welfare decomposition that separates trade, MP, and migration contributions—a result that does not appear in any of the predecessor frameworks.

There are N countries and J sectors; all markets are perfectly competitive, and labor is the sole factor of production. Table 4 summarizes notation; we adopt the {origin i, production l, destination n} convention throughout.

3.1 Preferences

Each country n is endowed with L_n workers. Workers consume composite goods from all J sectors with Cobb-Douglas preferences and expenditure shares α_n^j, $\sum_j \alpha_n^j = 1$. The aggregate price index in country n is $P_n = \prod_j (P_n^j/\alpha_n^j)^{\alpha_n^j}$, where P_n^j is the sector-j price index derived below. Real income I_n/P_n is the welfare measure for the representative agent; for Korea, where we model heterogeneous workers, we use the group-specific welfare W_{ig} defined in Section 3.7.

3.2 Technology

3.2.1 Composite intermediate goods

There is a continuum of tradable intermediate goods $v^j \in [0, 1]$ in each sector j. In country n, sector j composite intermediate goods are produced by sourcing intermediate goods v^j from the lowest-cost suppliers across countries and combining them through a CES production function with elasticity of substitution $\sigma^j > 0$. The production function for the composite intermediate good is given by

$$Q_n^j = \left[\int_0^1 q_n^j(v^j)^{(\sigma^j-1)/\sigma^j} dv^j\right]^{\sigma^j/(\sigma^j-1)} \tag{6}$$

Table 4: Summary of main notation

Indices and sets	
i	Country where the firm is headquartered;
l	Country where production takes place
n	Country where the product is consumed
j, k	Sector
Key variables	
w_i^j	Wage per efficiency unit in sector j of country i
P_n^j	CES price index for sector j in destination n
c_{il}^j	Unit production cost for a firm from i operating in location l, sector j
X_{iln}^j	Sales of sector-j output produced in l by an i-owned firm and consumed in n
L_r	Labor employed in region r
Friction parameters	
d_{ln}^j	Iceberg trade cost for shipping sector j from location l to destination n
h_{il}^j	Efficiency loss (cost wedge) when an i-owned plant produces sector j in l (MP cost)
ν_{rg}	Utility-reducing migration cost for a worker born in g who resides in r
Preference and technology parameters	
σ^j	Elasticity of substitution across varieties in sector j (CES demand aggregator)
ξ	Elasticity of substitution between home and host input bundles in multinational production ($\xi \neq \sigma^j$)
α_n^j	Final-expenditure (consumption) share of sector j for country n; $\sum_j \alpha_n^j = 1$
γ_i^j	Labor cost share in the Cobb-Douglas production function for sector j in country i
a^j	Share of home-sourced intermediates in multinational production of sector j

where $q_n^j(v^j)$ is the quantity of variety v^j used. Standard CES demand yields the price index

$$P_n^j = \left[\int p_n^j(v^j)^{1-\sigma^j} dv^j\right]^{\frac{1}{1-\sigma^j}} \tag{7}$$

The composite Q_n^j is used in three ways: as final consumption in country n; as materials for domestic and multinational production of intermediate good v^k by country n's firms in industry k; and as materials for domestic production of intermediate good v^k by foreign affiliates operating within country n. The specific usage of Q_n^j in MP is detailed in the following section.

3.2.2 Intermediate goods

Each country i can produce sector-j varieties at home or abroad. Country i's productivity for variety v^j across locations is $\mathbf{z}_i^j(v^j) \equiv \left\{z_{i1}^j(v^j), \ldots, z_{iN}^j(v^j)\right\}$, where $z_{il}^j(v^j)$ is productivity when producing in country l.

Production at home ($l = i$) is domestic production; production abroad ($l \neq i$) is multinational production (MP) by home country i in host l. Country i can serve destination $n \neq i$ in three ways: domestic production and export ($l = i$), horizontal MP ($l = n$), or bridge/export-platform MP ($l \neq i, l \neq n$).

Home production When the intermediate good is produced in home country i, it uses two inputs: labor and composite intermediate goods from all sectors, which are combined according to Cobb-Douglas production functions. The production function for the intermediate good is given by

$$q_i^j\left(v^j\right) = z_i^j\left(v^j\right)\left[H_i^j\left(v^j\right)\right]^{\gamma_i^j} \prod_{k=1}^{J}\left[m_i^{kj}\left(v^j\right)\right]^{\gamma_i^{kj}} \tag{8}$$

where $z_i^j\left(v^j\right) \equiv z_{ii}^j\left(v^j\right)$ represents country i's productivity in producing intermediate goods at home, $H_i^j\left(v^j\right)$ is the efficiency unit of labor used in the production of intermediate good v^j, and $m_i^{kj}\left(v^j\right)$ is the composite intermediate goods from sector k used for the production of intermediate good v^j. The Cobb-Douglas production function parameter $\gamma_i^{kj} \geqslant 0$ is the share of materials from sector k used in the production of intermediate good v^j, with $\sum_{k=1}^{J} \gamma_i^{kj} = 1 - \gamma_i^j$, where γ_i^j is the labor share of production.

The cost of the input bundle for producing intermediate goods at home, denoted by c_i^j, is then given by

$$c_i^j = \Upsilon_i^j (w_i^j)^{\gamma_i^j} \prod_{k=1}^{J} (P_i^k)^{\gamma_i^{kj}} \tag{9}$$

where w_i^j is the wage per efficiency unit in sector j of country i, P_i^k is the price of a composite intermediate good of sector k, and $\Upsilon_i^j \equiv (\gamma_i^j)^{\gamma_i^j} \prod_{k=1}^{J} (\gamma_i^{kj})^{\gamma_i^{kj}}$.

Since intermediate goods are produced with constant returns to scale technology, and all markets are perfectly competitive, firms price at unit cost. The unit cost of an intermediate good in country i produced at home is given by $c_i^j / z_i^j \left(v^j\right)$.

International trade costs An iceberg-type trade cost is denoted by $d_{in}^j \geq 1$. This cost is associated with shipping a sector j good from country i to country n. Specifically, delivering one unit of the sector j good from i to n requires d_{in}^j units to be shipped from i. We assume $d_{nn}^j = 1$ for all n and j and $d_{in}^j \leq d_{il}^j d_{ln}^j$ for all n, l, i and j.

Thus, the price of variety v^j exported from country i to n is $d_{in}^j c_i^j / z_i^j \left(v^j\right)$.

Multinational production costs Following Ramondo and Rodríguez-Clare (2013), MP involves two frictions. First, an iceberg-type efficiency loss $h_{il}^j \geq 1$ captures the cost of operating country i's technology in host l. Second, affiliates must source a fraction of their inputs from headquarters, combining home and host input bundles via CES technology. The production function of country i's affiliate producing v^j in host l is

$$q_{il}^j(v^j) = z_{il}^j(v^j) \left[(1 - a^j)^{\frac{1}{\xi}} \left(M_l^j(v^j) \right)^{\frac{\xi-1}{\xi}} + (a^j)^{\frac{1}{\xi}} \left(M_i^j(v^j) \right)^{\frac{\xi-1}{\xi}} \right]^{\frac{\xi}{\xi-1}} \tag{10}$$

where M_l^j and M_i^j are input bundles from host l and home i, respectively. The parameter $a^j \in (0,1)$ governs the strength of the headquarters linkage: a larger a^j means affiliates rely more heavily on home-sourced inputs, creating the vertical complementarity documented in Section 2.4—as Korean firms expand production in China, they pull intermediates from domestic parents, generating positive employment effects at home. The elasticity $\xi > 1$ controls how readily affiliates substitute between home and host inputs as relative prices change.

Home inputs shipped to the host incur trade costs, so the effective price of the home bundle is $c_i^j d_{il}^j$; the host bundle costs $c_l^j h_{il}^j$ after the efficiency loss. The affiliate's unit cost c_{il}^j is therefore

$$c_{il}^j = \left[(1 - a^j) \left(c_l^j h_{il}^j \right)^{1-\xi} + a^j \left(c_i^j d_{il}^j \right)^{1-\xi} \right]^{1/(1-\xi)} \tag{11}$$

where c_l^j is defined in equation (9), and $c_{ii}^j = c_i^j$ by construction. The parameter a^j is estimated from OECD AAMNE-XVEM data on within-affiliate input sourcing patterns following Alviarez (2019); our estimates yield $a^j \in [0.04, 0.08]$ across tradeable sectors.

In country l, the unit cost of an intermediate good produced in country l by foreign affiliates from country $i \neq l$ is given by $c_{il}^j / z_{il}^j(v^j)$.

Productivity distribution Following Eaton and Kortum (2002), the productivity vectors $\mathbf{z}_i^j(v^j)$ are assumed to be independently drawn across countries, sectors, and goods from the following multivariate Fréchet distribution:

$$F_i^j(\mathbf{z}_i^j) = \exp\left\{ -T_i^j \left[\sum_l \left(z_{il}^j \right)^{-\theta^j/(1-\rho)} \right]^{1-\rho} \right\} \tag{12}$$

where T_i^j is a country- and sector-specific scale parameter, θ^j is sector-specific shape parameter with $\theta^j > \max\{1, \sigma^j - 1\}$,[12] and ρ determines the degree of correlation among the elements of $\mathbf{z}_i^j$. When ρ is high, a higher productivity draw for producing in one country is more strongly associated with a higher productivity draw for producing in another country. We calibrate $\rho = 0.5$ following Ramondo and Rodríguez-Clare (2013), implying moderate positive correlation: a firm that is highly productive domestically tends to be above-average abroad. This generates the complementarity between home and overseas production that underlies the positive $\hat{\beta}_{MP}$ estimated in Section 2.4: firms that expand affiliate production in China are disproportionately those with high domestic productivity, so their expansion abroad coexists with—rather than displaces—strong home activity.

[12] The condition $\theta^j > \sigma^j - 1$ ensures the CES price index integral converges (the Gamma function $\Gamma(1-(\sigma^j-1)/\theta^j)$ is well-defined), and $\theta^j > 1$ ensures the Fréchet distribution has a finite mean so expected productivity is well-defined. We use sector-specific σ^j from Broda and Weinstein (2006) (see Section 4), which satisfies this condition for all ten sectors; the tightest margins are Food ($\theta^{\text{Food}} = 2.62 > \sigma^{\text{Food}} - 1 = 2.0$) and Chemicals ($3.13 > 2.5$).

3.3 Prices and expenditure shares

Taking international trade costs into account, the unit cost of a sector j intermediate good in country n, produced in country l with a technology from country i is given by $c_{il}^j d_{ln}^j / z_{il}^j(v^j)$. Therefore, the price of intermediate good v^j in country n is given by

$$p_n^j(v^j) = \min_{i,l} \frac{c_{il}^j d_{ln}^j}{z_{il}^j(v^j)}$$

By the assumption of a multivariate Fréchet distribution for productivity draws, as specified in equation (12), the share of expenditures by country n on intermediate goods produced in country l with technologies from country i is given by

$$\pi_{iln}^j = \lambda_{in}^j \lambda_{iln}^j = \frac{T_i^j \left(\tilde{c}_{in}^j\right)^{-\theta^j}}{\sum_k T_k^j \left(\tilde{c}_{kn}^j\right)^{-\theta^j}} \frac{\left(c_{il}^j d_{ln}^j\right)^{-\theta^j/(1-\rho)}}{\sum_k \left(c_{ik}^j d_{kn}^j\right)^{-\theta^j/(1-\rho)}} \tag{13}$$

where $\tilde{c}_{in}^j \equiv \left[\sum_k \left(c_{ik}^j d_{kn}^j\right)^{-\theta^j/(1-\rho)}\right]^{-(1-\rho)/\theta^j}$. The first term on the right-hand side, denoted by λ_{in}^j, represents the share of expenditures by country n on intermediate goods produced with technologies from country i. The second term, denoted by λ_{iln}^j, indicates the share of these goods produced in country l.

The price index of sector j composite intermediate goods in country n is given by

$$P_n^j = \Gamma^j \left[\sum_i T_i^j \left(\tilde{c}_{in}^j\right)^{-\theta^j}\right]^{-1/\theta^j} \tag{14}$$

where $\Gamma^j = \left[\Gamma\left(\frac{\theta^j+1-\sigma^j}{\theta^j}\right)\right]^{\frac{1}{1-\sigma^j}}$

3.4 Goods market clearing

Let Y_{il}^j be the value of sector j intermediate goods produced in country l with technologies from country i. Then, Y_{il}^j is given by

$$Y_{il}^j = \sum_n \pi_{iln}^j P_n^j Q_n^j \tag{15}$$

where Q_n^j and P_n^j are defined in equations (6) and (7), $P_n^j Q_n^j$ are the values of sector j composite input produced in country n.

$P_n^j Q_n^j$ is allocated across country n's domestic production, multinational production in $i \neq n$ and domestic consumption. Specifically, $P_n^j Q_n^j$ is given by

$$P_n^j Q_n^j = \sum_k \gamma_n^{jk} Y_{nn}^k + \sum_k \gamma_n^{jk} \sum_{i \neq n} (1 - \omega_{in}^k) Y_{in}^k + \sum_k \gamma_n^{jk} \sum_{l \neq n} \omega_{nl}^k Y_{nl}^k + \alpha_n^j I_n \tag{16}$$

where $\omega_{in}^k = a^k \left(c_i^k d_{in}^k / c_{in}^k \right)^{1-\xi}$ is the cost share of home-sourced inputs for affiliates from i producing in n, and $I_n = \sum_g I_{ng} + D_n$ is aggregate income (labor income plus trade deficit). The four terms on the right-hand side decompose expenditure by source:

(i) $\sum_k \gamma_n^{jk} Y_{nn}^k$: intermediate demand from country n's domestic production;

(ii) $\sum_k \gamma_n^{jk} \sum_{i \neq n} (1 - \omega_{in}^k) Y_{in}^k$: host-sourced input demand by foreign affiliates operating in n;

(iii) $\sum_k \gamma_n^{jk} \sum_{l \neq n} \omega_{nl}^k Y_{nl}^k$: home-sourced input demand by n's own affiliates operating abroad—these are the intra-firm shipments that generate the vertical complementarity channel;

(iv) $\alpha_n^j I_n$: final consumption.

Terms (ii) and (iii) capture intra-firm trade in inputs, connecting to the empirical moment used to calibrate a^j (Section 4).

3.5 Trade flows

Let M_{ln}^j denote imports of sector-j goods by country n from $l \neq n$. Total imports comprise arm's-length trade in intermediate goods and intra-firm shipments from l's affiliates operating in n. This decomposition is important because the empirical trade data used in Section 2.4 measure gross bilateral flows, which bundle both components; the model makes the distinction explicit:

$$M_{ln}^j = \sum_i \pi_{iln}^j P_n^j Q_n^j + \sum_k \gamma_l^{jk} \omega_{ln}^k Y_{ln}^k + \gamma_l^j \omega_{ln}^j Y_{ln}^j \tag{17}$$

where ω_{ln}^j is defined as in equation (16).

Total imports by country n from $l \neq n$ is given by

$$M_{ln} = \sum_j M_{ln}^j \tag{18}$$

The trade balance condition is given by

$$D_n = \sum_{l \neq n} M_{ln} - \sum_{l \neq n} M_{nl} \tag{19}$$

where D_n is exogenous in the model.

3.6 Labor demand

Let LD_i^j denote the demand for efficiency units of labor in sector j of country i. Analogous to the goods market clearing condition, the total demand for efficiency units of labor comprises three components: demand from country i's domestic production, demand from foreign countries producing in i, and demand from country i when producing abroad.

$$LD_i^j \equiv \frac{\gamma_i^j}{w_i^j} \left(Y_{ii}^j + \sum_{n \neq i} Y_{ni}^j (1 - \omega_{ni}^j) + \sum_{l \neq i} Y_{il}^j \omega_{il}^j \right) \tag{20}$$

where γ_i^j is the share of value added in producing intermediate goods at home, and w_i^j is the wage per efficiency unit in sector j of country i.

3.7 Labor supply

Labor supply is determined by workers' migration and sector choices, following Galle et al. (2023). This is the channel through which the China shock generates cross-regional welfare inequality: migration frictions prevent workers from fully reallocating in response to sector-specific shocks, so initial location determines exposure—the mechanism that the SUTVA discussion in Section 2.4 identified as requiring a GE treatment.

Each country n is endowed with a measure of L_n workers, which are mobile across sectors and regions within a country. There are G_n groups of workers in country n, and we define groups based on geographic regions. Specifically, we set $G_{Korea} = 5, G_n = 1$ for all $n \neq$ Korea. Let G_n denote the set of regions in country n as well. Workers within a group $g \in G_{Korea}$ are referred to as "from region g." Each region g in country n is endowed with a measure of L_{ng} workers, where $\sum_{g \in G_n} L_{ng} = L_n$. We use g, h, and f to denote the regions within a country.

In country i, each worker has a number of efficiency unit $\boldsymbol{b} = \left\{ b_h^j \right\}_{h \in G_i, j \in J}$ in each region-sector combination (h, j) drawn from a Fréchet distribution with shape parameter κ_g—which depends on the worker's region of origin g—and scale parameters A^j_{ih}. A^j_{ih} collects all working region-sector factors that scale the attractiveness of h (for example, skill–specific amenity bundles or baseline comparative advantages). Workers are fully characterized by their efficiency units $\boldsymbol{b}$ and an origin region g. We require $\kappa_g > 1$ for all g so that the expected wage—which involves $\Gamma(1 - 1/\kappa_g)$—is finite and well-defined.

Let w^j_i denote wages per efficiency unit in sector j of country i. When a worker from region g in country i works in region h of country i, they experience a proportional adjustment to income, represented by ν_{igh}, with $\nu_{igg} = 1$ and $\nu_{igh} \leq 1$ for all i, g, h. Then, a worker from g that works in region h in sector j earns income of $w^j_i \nu_{igh} b^j_h$.

Let Ω^j_{igf} denote the measure of workers from region g who choose sector j and region f:

$$\Omega^j_{igf} \equiv \left\{ \boldsymbol{b} \text{ s.t. } w^j_i \nu_{igf} b^j_f \geq w^k_i \nu_{igh} b^k_h \text{ for all } h, k \right\}.$$

A worker with $\boldsymbol{b}$ from region g in country i will choose to work in region-sector (f, j) if and only if $\boldsymbol{b} \in \Omega^j_{igf}$. The share of workers from region g in country i that choose to work in (f, j) is

$$\mu^j_{igf} \equiv \int_{\Omega^j_{igf}} dF(\boldsymbol{b}) = \frac{A^j_{if} \left(\nu_{igf} w^j_i \right)^{\kappa_g}}{(\Phi_{ig})^{\kappa_g}} \tag{21}$$

where $\Phi_{ig} \equiv (\sum_{h,k} A^k_{ih} \left(\nu_{igh} w^k_i \right)^{\kappa_g})^{1/\kappa_g}$.

Two forces govern relocation incentives in this expression. First, higher wages raise the numerator in the share expression and, all else equal, increase the share of workers who settle in region f. Second, relocation friction $\nu_{igf} \leq 1$—which can reflect geographic distance, housing prices, or foregone social networks—dampens that effective reward in region f. The parameter κ_g controls the sensitivity of labor flows to these differences for workers from region g: when κ_g is high, small gaps in the log wage–net–of–cost translate into pronounced shifts in residency shares, whereas a low κ_g implies a flatter response and thus more muted migration across regions. Because κ_g is region-of-origin specific, workers from different regions respond heterogeneously to the same wage incentives, generating cross-regional heterogeneity in welfare gains even when wage changes are uniform across groups.

The efficiency units supplied by the group g in region-sector (f, j) are given by

$$\mathcal{Z}^j_{igf} \equiv L_{ig} \int_{\Omega^j_{igf}} b^j_f dF_i(b) = \mu^j_{igf} \eta_g L_{ig} \frac{\Phi_{ig}}{w^j_i \nu_{igf}} \tag{22}$$

where $\eta_g \equiv \Gamma(1 - 1/\kappa_g)$, and L_{ig} is the number of workers from region g.

Total income of group g in country i is $I_{ig} \equiv \sum_{f,j} w^j_i \nu_{igf} \mathcal{Z}^j_{igf} = \eta_g L_{ig} \Phi_{ig}$. The share of income obtained by workers in group g in country i in region-sector (f, j) is also given by μ^j_{igf}.

We define the measure of the welfare of workers from region g in the country i as their ex-ante real income.

$$W_{ig} \equiv \frac{I_{ig}/L_{ig}}{P_i} \tag{23}$$

where $P_i = \prod_j \left(\frac{P^j_i}{\alpha^j_i}\right)^{\alpha^j_i}$.

The excess demand for efficiency units in sector j of country i is

$$ELD^j_i \equiv LD^j_i - \sum_{f \in G_i} \sum_{g \in G_i} \mathcal{Z}^j_{igf} \tag{24}$$

and in equilibrium, the following conditions hold for all i and j:

$$ELD^j_i = 0 \tag{25}$$

Definition 1 *Given* $\{L_{ig}\}, \{d^j_{in}\}, \{h^j_{in}\}, \{T^j_i\}, \{A^j_{ig}\}, \{D_i\}$*, an equilibrium is a vector of wages* $\{w^j_i\}_{i \in N, j \in J}$ *and prices* $\{P^j_i\}_{i \in N, j \in J}$ *that satisfies equilibrium conditions (9), (11), (13), (14), (15), (16), (21), (22), (20) and (25). By Walras' law, one of the* $2NJ$ *equilibrium conditions is redundant, so wages and prices are determined only up to a common normalization. We fix the composite wage index of the United States to unity,* $\bar{w}_{USA} \equiv \sum_j \alpha^j_{USA} w^j_{USA} = 1$*, which makes all wages and prices commensurate with US efficiency-unit labor as the numéraire. Welfare counterfactuals are expressed as hat-algebra ratios and are invariant to this choice.*[13]

[13] The system of $2NJ$ nonlinear equations does not, in general, have a unique solution. Models with input-output linkages and endogenous migration can admit multiple equilibria (Allen and Arkolakis, 2014). We follow the standard approach in the quantitative trade literature: we solve for the equilibrium using iterative methods (Appendix E) initialized at the observed data, and verify that the algorithm converges to the same fixed point from multiple starting values. In the hat-algebra counterfactuals, uniqueness of the *change* in equilibrium is more likely

All equilibrium objects can be expressed as functions of wages $\{w_i^j\}$ and prices $\{P_i^j\}$: unit costs via (9)–(11), expenditure shares via (13), and quantities via (16). The $2NJ$ conditions from (14) and (25) then determine the equilibrium. Appendix E describes the nested fixed-point algorithm used to solve this system.

Consider shocks to trade costs or MP costs. We can derive a welfare accounting identity in the spirit of Arkolakis et al. (2012), extended here to incorporate migration reallocation and multinational production channels that do not appear in their original single-country framework. Proposition 1 below is therefore a distinct result: Arkolakis et al. (2012) applies to a representative-agent one-country model with fixed labor supply, whereas Proposition 1 accommodates heterogeneous workers, multiple regions, and endogenous migration, so the welfare formula includes an additional migration term ($\hat{\mu}_{igg}^k$) that has no counterpart in their sufficient-statistic result. Define the notation as $\hat{x} \equiv x'/x$. Let $\lambda_{iii}^k \equiv \pi_{iii}^k/\lambda_{ii}^k$ denote the share of country i's technology goods consumed in i that are also *produced* in i (as opposed to produced by i's foreign affiliates and re-exported back). This is the MP-openness term: a fall in λ_{iii}^k reflects an expansion of affiliate production abroad, reducing the domestic-production share.

Proposition 1 (Welfare Decomposition) *Given some shocks to trade costs, MP costs, or both, the percentage change in the real wage of group g in country i is given by*

$$\hat{W}_{ig} = \prod_{j,k} (\hat{\lambda}_{ii}^k)^{-\alpha_i^j \tilde{a}_i^{jk}/\theta^k} \prod_{j,k} (\hat{\lambda}_{iii}^k)^{-(1-\rho)\alpha_i^j \tilde{a}_i^{jk}/\theta^k} \prod_{j,k} (\hat{\mu}_{igg}^k)^{-\alpha_i^j \tilde{a}_i^{jk} \gamma_i^k/\kappa_g} \tag{26}$$

where $\tilde{a}_i^{jk}$ is the typical element of matrix $\left(I - \boldsymbol{\gamma}_i^T\right)^{-1}$ with $\boldsymbol{\gamma}_i \equiv \left\{\gamma_i^{kj}\right\}_{k\in J, j\in J}$ the $J \times J$ matrix of IO coefficients, and κ_g is the region-of-origin-specific migration elasticity estimated by MLE.

The first and second terms on the right-hand side of (26) represent common shocks to all groups in country i. $\hat{\lambda}_{ii}^k$ captures changes in the home-technology expenditure share (a trade-openness term), $\hat{\lambda}_{iii}^k$ captures changes in the domestic-production share of home-technology goods (an MP-openness term), and $\hat{\mu}_{igg}^k$ reflects how workers in group g adjust their sectoral and regional employment in response. Therefore, the welfare gains of group g depend on both changes in openness and the group's labor adjustment.

to hold, since the system is solved as a perturbation around the observed equilibrium rather than from primitives (Dekle et al., 2007).

Remark (role of a^j in the MP term). The MP channel in Proposition 1 operates through two distinct mechanisms. First, $\hat{\lambda}^k_{iii}$ responds to MP cost changes h^j_{il} regardless of a^j: lower MP costs shift production from home to affiliates, reducing the domestic-production share. Second, the *vertical fragmentation channel*—the feedback of MP costs through affiliate marginal costs via $\omega^j_{ln} = a^j(\cdot)^{1-\xi}$—is active only when $a^j > 0$. With our estimated $a^j \in [0.04, 0.08]$ (Section 4), both channels are operative; Section 5 quantifies their relative contributions. Appendix C compares the model's structure with Caliendo and Parro (2015), Antràs and De Gortari (2020), and De Gortari (2019), and discusses the assumption of uniform within-sector wages across Korean regions.

Summary. The model delivers three objects that the reduced-form analysis in Section 2 cannot provide on its own. First, a welfare decomposition (Proposition 1) that separately attributes gains to trade openness ($\hat{\lambda}^k_{ii}$), MP openness ($\hat{\lambda}^k_{iii}$), and labor reallocation ($\hat{\mu}^k_{igg}$). Second, general-equilibrium counterfactuals that account for cross-regional spillovers through migration and input-output linkages—the SUTVA violation identified in Section 2.4. Third, a framework for comparing trade-only and trade-plus-MP models, quantifying the welfare mismeasurement from omitting the MP channel. We turn next to parameterization.

4 Parameterization

We parameterize the model using data around the year 2000, where available. Our analysis covers ten countries—Canada, China, Germany, Japan, Mexico, South Korea, Taiwan, the United Kingdom, and the United States—along with the rest of the world (ROW). For Korea, we further distinguish five regions. We consider 10 sectors following Alviarez (2019): Food, Textiles, Wood and Paper, Chemicals, Minerals, Electrical and Machinery, Metals, Transport, Miscellaneous Manufacturing, and Service.

We classify the parameters into two groups. The first group consists of externally calibrated parameters, either directly set to match the data or assigned values based on estimations or calibrations in the literature. The second group includes parameters whose values are chosen to ensure that the model's equilibrium outcomes align with observed data.

4.1 Parameters calibrated externally

Consumption and IO shares. We set the final consumption share α_n^j equal to the sectoral share of final consumption expenditure by households, obtained from the OECD Input-Output tables for each country. For the ROW, we use data covering 57 countries. We set the IO coefficients for domestic production, $\gamma_n^j, \gamma_n^{j,k}$ based on the OECD Inter-Country Input-Output tables, which are disaggregated by ownership (domestic-owned and foreign-owned firms[14]). For the ROW, we include 66 countries along with the aggregated rest-of-world category.

Labor endowments. Labor endowments, $L_n, L_{Korea,g}$ are set to match the number of total employment obtained from the Penn World Table and the Korea Labor & Income Panel Study, respectively.

Substitution elasticity (ξ) and productivity correlation (ρ). The elasticity of substitution between the home input bundle and the host input bundle, ξ is set to 1.5, following Ramondo and Rodríguez-Clare (2013). This choice aligns with the cross-wage elasticities of labor demand for German multinationals across countries, as estimated by Muendler and Becker (2010). For the

[14] In this OECD database, firms are categorized based on ownership structure. Foreign-owned firms are defined as firms in which at least 50% of ownership is held by foreign entities.

correlation parameter of the multivariate Fréchet distribution of productivity, we follow Ramondo and Rodríguez-Clare (2013) and set $\rho = 0.5$, since this parameter is not well identified using aggregate data on trade and multinational production. Table 12 in Section 5 shows that the welfare gains are robust to setting $\rho = 0$.

Trade elasticity (θ^j). The shape parameter of the productivity distribution, θ^j, is set to sector-specific values from Caliendo and Parro (2015) Table 1 (99% sample benchmark): $\theta^j \in \{2.62, 8.10, 11.50, 3.13, 13.53,$ for sectors Food, Textiles, Wood, Chemicals, Mining, Electrical Machinery, Metal Products, Transport, Miscellaneous, and Nontradables, respectively.[15] These estimates use triple-differenced gravity equations and are the standard benchmark for quantitative trade models with Fréchet-distributed productivity.

Remark (effective trade elasticity). In our model, the effective trade elasticity is $\theta^j/(1-\rho)$ rather than θ^j alone, because the multivariate Fréchet distribution with $\rho > 0$ compresses the dispersion of the delivered-cost distribution (Ramondo and Rodríguez-Clare, 2013). With $\rho = 0.5$, the effective elasticity is $2\theta^j$, so the Caliendo-Parro estimates implicitly understate the model's responsiveness to trade costs. We retain their θ^j values for comparability; Table 14 in Section 5 quantifies the sensitivity of welfare results to this choice.

Migration elasticity (κ). The shape parameter for the distribution of workers' productivity, κ, is set to 1.5 using the value from Galle et al. (2023) for India as a benchmark, consistent with our own Korean MLE estimates (Table 5, cross-region average ≈ 1.54).

4.2 Parameters calibrated in equilibrium

Six parameter groups are calibrated within the model equilibrium, each pinned to a distinct data moment: the gravity coefficients Δ target bilateral trade and MP shares (R^{Trade}, R^{MP}); the home-input shares a^j match intrafirm import ratios of foreign affiliates in Korea; the productivity locations T_i^j match sectoral real wages; the migration costs ν_{igf} match inter-regional migration flows; and the worker-productivity locations A_{if}^j match regional employment shares by sector. The

[15] Sectors with sign changes in Caliendo and Parro (2015) (Basic Metals, Machinery n.e.c., Motor Vehicles, Other Transport) are assigned the cross-sector mean of 4.45, following their paper's methodology. The Nontradables sector is assigned the same value as a placeholder; θ^j is irrelevant for this sector since $d_{in}^j = \infty$ for $j =$ Nontradables.

remainder of this subsection describes each step and its identifying variation.

We assume that bilateral trade and MP costs are a function of distance between two countries and whether countries share a border or a language:

$$d_{in}^{j} = \exp(\delta_0^j + \delta_{dist}^j dist_{in} + \delta_{bord}^j b_{in} + \delta_{lang}^j l_{in})$$

$$h_{in}^{j} = \exp(\zeta_0^j + \zeta_{dist}^j dist_{in} + \zeta_{bord}^j b_{in} + \zeta_{lang}^j l_{in})$$

for all $i \neq n$, with $d_{ii}^j = 1$ and $h_{ii}^j = 1$ for all j.

Let $\Delta \equiv \{\delta_0^j, \delta_{dist}^j, \delta_{bord}^j, \delta_{lang}^j, \zeta_0^j, \zeta_{dist}^j, \zeta_{bord}^j, \zeta_{lang}^j\}$. The full parameter vector to calibrate is $[\Delta, a^j, A_{if}^j, \nu_{igf}, T_i^j, d_{in}^j, h_{in}^j]$.

Given Δ, bilateral costs d_{in}^j and h_{in}^j follow immediately. The home-input shares a^j are then chosen to match the model's expenditure share on the home input bundle, $\omega_{il}^j = a^j(c_i^j d_{il}^j / c_{il}^j)^{1-\xi}$, averaged across source countries i for affiliates operating in Korea, to its data counterpart: the ratio of intrafirm imports by foreign affiliates in Korea to their Korean sales, averaged across countries within each sector. The resulting estimates are $a^j \in [0.04, 0.08]$ for tradeable sectors and $a^j = 0$ for nontradables; all counterfactual exercises in Section 5 are run with these estimated values.[16]

The productivity locations T_i^j are calibrated to match sectoral real wages w_i^j / P_i.

Given $[\Delta, a^j, T_i^j]$, we calibrate migration costs in Korea, $\nu_{Korea,gf}$, and the worker-productivity locations $A_{i,f}$. Migration costs are parameterized as $\ln \nu_{Korea,gf} = -\beta_0 - \beta_1 \cdot dist_{gf}$, where $dist_{gf}$ is the distance between the City Hall or Provincial Government Office of each region's central city. In the outer loop, we select β_0, β_1 to minimize the sum of the squared deviations between migration shares in the model, μ_{igf}^j, and their data counterparts. In the inner loop, we calibrate $A_{i,f}$ such that the sectoral employment share in each region, $\frac{\sum_g L_{i,g} \mu_{igf}^j}{L_i}$, matches the corresponding values in the data. For ROW countries ($i \neq$ Korea), the location parameters $A_{i,f}^j$ are calibrated to match country-level occupation shares $\mu_{ROW,i}^j$ using a simultaneous proportional updating algorithm.[17]

[16] Specifically, the nine tradeable-sector estimates are: Food 0.035, Textiles 0.065, Wood 0.058, Chemicals 0.076, Minerals 0.036, Electrical/Machinery 0.084, Metals 0.084, Transport 0.080, Miscellaneous 0.054. These values are stored in `fundamentals.a` in the calibrated equilibrium file and are active in all counterfactual computations.

[17] In the preliminary implementation, Korean regional employment shares are matched to within 1.2% (max absolute deviation across regions and sectors) and ROW occupation shares to within 12%. The 12% ROW tolerance reflects the best achievable fit of the current model structure: with $N = 10$ countries, $J = 10$ sectors, and aggregate OECD ICIO employment data, the model's equilibrium occupation shares cannot exactly replicate the observed country-level distribution. Diagnostic tests confirm this is a structural limitation rather than an algorithmic failure—

The equilibrium wage of a worker from region g is

$$\tilde{w}_{ig} \equiv \max_{f \in G_i,\, j \in J} \; w_i^j \, \nu_{igf} \, b_f^j.$$

Then its CDF is

$$G_{ig}(w) \equiv \Pr(\tilde{w}_{ig} \leq w) = \exp\left[-\left(\sum_{f \in G_i} \sum_{j \in J} A_{if}^j \, (w_i^j \nu_{igf})^{\kappa_g}\right) w^{-\kappa_g}\right],$$

which is a Fréchet distribution with shape parameter κ_g and scale parameter $\sum_{f,j} A_{if}^j \, (w_i^j \nu_{igf})^{\kappa_g}$. This closed-form CDF is the basis for the maximum likelihood estimation of κ_g reported in Table 5: we observe the realized wage distribution across workers from each origin region g and maximize the Fréchet log-likelihood over κ_g.

Table 5: Maximum Likelihood Estimates of Labor Allocation Elasticity

	Worker Origin Region				
Parameter	Seoul	Northwest	Northeast	Southwest	Southeast
κ	1.91	1.49	1.48	1.28	1.54
	(0.0442)	(0.0220)	(0.0483)	(0.0213)	(0.0207)

Note: Standard errors are in parentheses.

Table 5 reports the maximum likelihood estimates of κ by worker origin region. The estimates range from 1.28 (Southwest) to 1.91 (Seoul), with a cross-region average of approximately 1.54, closely comparable to $\kappa = 1.5$ from Galle et al. (2023) for India. The regional variation implies that workers in some regions are more homogeneous in their skills across sectors (high κ, more elastic reallocation) than others, which would affect both the magnitude and the regional distribution of welfare responses to the China shock. The counterfactual analyses in Section 5 use the uniform value $\kappa = 1.5$ as the baseline; Table 13 confirms that the welfare decomposition is robust to varying κ across the range implied by our MLE estimates.

reducing the step size below 0.01 causes the ROW error to increase rather than decrease, consistent with a fixed point of the proportional updating map at approximately 12% max deviation. The Korea calibration converges cleanly to 1.2%, suggesting the aggregate ROW approximation is the binding constraint. To assess sensitivity, we re-solve the model replacing ROW $A_{i,f}^j$ with their upper- and lower-bound values consistent with the 12% tolerance; the resulting change in Korea's aggregate welfare estimate is less than 0.05 percentage points, confirming that the ROW miscalibration does not materially affect the Korean results reported in Section 5.

$$\nu_{igf} = \left(\frac{\mu_{igf}\mu_{ifg}}{\mu_{igg}\mu_{iff}} \right)^{\frac{1}{2\kappa}}$$

where $\mu_{igf} = \sum_j \mu_{igf}^j$.[18]

Given the estimated Φ_{ig}^{κ} and observed migration shares μ_{igf}^j, we back out A_{if}^j up to normalization (setting $A_1^j = 1$) as in Lee and Yi (2018):

$$A_f^j = \frac{\bar{A}_{gf}^j}{\bar{A}_{g1}^j} \left(\frac{\nu_{g1}}{\nu_{gf}} \right)^{\kappa},$$

where $\bar{A}_{gf}^j \equiv (\mu_{gf}^j)^{\text{data}} (\Phi_{ig}^{\kappa})^{\text{est}}$. The full derivation is in Appendix D. Migration share data are obtained from the Korean Labor & Income Panel Study; employment shares across sectors and regions are sourced from the Korean Mining and Manufacturing Survey and the Korean Labor Force Survey.

Then, given $[\Delta, a^j, A_{if}^j, \nu_{igf}, T_i^j, d_{in}^j, h_{in}^j]$, we solve for the equilibrium wages and prices. We compute the bilateral trade and MP shares, defined as $\lambda_{in}^{j,\text{Trade}} \equiv \frac{M_{in}^j}{\sum_l M_{ln}^j}$ and $\lambda_{in}^{j,\text{MP}} \equiv \frac{Y_{in}^j}{\sum_l Y_{ln}^j}$. Finally, we calculate a measure of the model's explanatory power for bilateral trade and MP shares as follows:

$$R^{Trade} \equiv 1 - \frac{\sum_{j,i,n} \left(\lambda_{in}^{j,Trade \text{ data}} - \lambda_{in}^{j,Trade, \text{ model}} \right)^2}{\sum_{j,i,n} \left(\lambda_{in}^{j,Trade, \text{ data}} \right)^2}$$

$$R^{MP} \equiv 1 - \frac{\sum_{j,i,n} \left(\lambda_{in}^{j,MP \text{ data}} - \lambda_{in}^{j,MP, \text{ model}} \right)^2}{\sum_{j,i,n} \left(\lambda_{in}^{j,MP, \text{ data}} \right)^2}$$

The parameters in Δ are chosen to minimize $(1-R^{Trade})+(1-R^{MP})$. In the current calibration, $R^{Trade} = 0.90$ and $R^{MP} = 0.79$, indicating that the model matches 90% of the variation in domestic trade shares and 79% of the variation in domestic MP shares across country-sector pairs.

[18] This formula is derived from the migration share equations. From equation (21), the ratio of off-diagonal to diagonal shares is $\mu_{igf}^j/\mu_{igg}^j = (A_{if}^j/A_{ig}^j)\nu_{igf}^{\kappa}$ and $\mu_{ifg}^j/\mu_{iff}^j = (A_{ig}^j/A_{if}^j)\nu_{ifg}^{\kappa}$. Multiplying these two ratios, the productivity terms A_{ig}^j/A_{if}^j cancel, yielding $(\mu_{igf}^j \mu_{ifg}^j)/(\mu_{igg}^j \mu_{iff}^j) = (\nu_{igf}\nu_{ifg})^{\kappa}$. Under the symmetry assumption $\nu_{igf} = \nu_{ifg}$, taking the $1/(2\kappa)$ power and aggregating over sectors gives the stated formula. The aggregation uses $\mu_{igf} = \sum_j \mu_{igf}^j$ with the same symmetry imposed at the aggregate level.

Notation	Description	Source/Value
A. Parameters Calibrated Externally		
α_n^j	Final consumption share	OECD Input-Output tables
γ_n^j, $\gamma_n^{j,k}$	IO coefficients when producing at home	OECD Inter-Country Input-Output tables split according to ownership
$L_n, L_{Korea,g}$	Labor endowment	Penn World Table, Korean Labor & Income Panel Study
σ^j	Elasticity of substitution across varieties within sector j (CES demand)	Broda and Weinstein 2006: sector-specific (3.0–6.0)[§]
ξ	Elasticity of substitution across input bundles in MP ($\xi \neq \sigma^j$)	Ramondo and Rodríguez-Clare (2013): 1.5
ρ	Correlation parameter of the multivariate Fréchet distribution of productivity	Ramondo and Rodríguez-Clare (2013): 0.5
θ^j	Shape parameter of the multivariate Fréchet distribution of productivity	Caliendo and Parro (2015): sector-specific (2.62–13.53)[†]
κ_g	Shape parameter of the multivariate Fréchet distribution of worker's productivity	MLE on KLIPS panel: region-specific (1.28–1.91)[‡]
B. Parameters Calibrated in Equilibrium		
Notation	**Description**	**Target**
a^j	Share of home input bundle	Ratio of intrafirm trade to multinational production by multinational firms operating in Korea
A_{if}^j	Location parameter of the multivariate Fréchet distribution of worker's productivity	Employment share across regions and sectors
ν_{igf}	Migration costs	Migration flows in Korea
T_i^j	Location parameter of the multivariate Fréchet distribution of productivity	Real wages
d_{in}^j	Trade costs	Import shares
h_{in}^j	MP costs	MP shares

Table 6: Parameterization

§ Sector-specific, aggregated from Broda and Weinstein (2006) HS 3-digit median estimates to our 10-sector ISIC/ICIO classification (range: 3.0–6.0), replacing the uniform $\sigma^j = 4$ of Ramondo and Rodríguez-Clare (2013) (one-sector calibration). All values satisfy $\theta^j > \sigma^j - 1$ for convergence of the CES price-index integral; tightest margins are Food ($2.62 > 2.0$) and Chemicals ($3.13 > 2.5$). Full sector-by-sector values are reported in Appendix D.

† Sector-specific estimates from Caliendo and Parro (2015) Table 1 (99% sample), ranging from 2.62 (Food) to 13.53 (Mining). Obtained from triple-differenced gravity equations and standard in quantitative trade models.

‡ Estimated by maximum likelihood on the Korean Labor and Income Panel Study (KLIPS). Region-specific estimates (Table 5): Seoul 1.91, Northwest 1.49, Northeast 1.48, Southwest 1.28, Southeast 1.54. The welfare formula (Proposition 1) uses region-specific κ_g in the migration term.

Table 6 summarizes all parameters and their identifying variation. With the calibrated equilibrium in hand, we now turn to counterfactual experiments that quantify the trade, MP, and migration channels of the China shock for Korean regional welfare.

5 Counterfactuals

We conduct two counterfactual exercises. The *preferred exercise* compares two equilibria calibrated from OECD ICIO data: a 2000 baseline (pre-WTO accession) and a 2007 equilibrium (post-integration). The *autarky comparison* sets China's trade and MP costs to prohibitively high values, characterizing the total gains from openness. In both exercises, we compare the full model to a trade-only counterfactual ($h^j_{in} \to \infty$ for $i \neq n$), isolating the MP channel's contribution—the structural analog of the reduced-form omitted variable bias documented in Section 2.

5.1 The China Shock

We calibrate the China shock as the change in bilateral trade and MP costs between 2000 (pre-WTO accession) and 2007 (post-integration), following Caliendo and Parro (2015) and Galle et al. (2023). For each year, bilateral trade costs d^j_{in} and MP costs h^j_{in} are backed out from observed OECD AAMNE-ICIO trade and MP shares using the Novy (2013) structural gravity formula:[19]

$$d^j_{in} \;\propto\; \left(\frac{\pi^j_{ni}\,\pi^j_{in}}{\pi^j_{nn}\,\pi^j_{ii}}\right)^{-1/(2\theta^j)} .$$

The China shock is then the transition from the 2000 cost equilibrium to the 2007 cost equilibrium, holding all other fundamentals (A^j_{if}, T^j_i, L, κ, θ^j) fixed at their calibrated 2000 values. This two-equilibrium comparison isolates the trade and MP cost reduction components of China's WTO accession and subsequent integration.[20] As a robustness check, we also report results from an autarky-to-baseline comparison ($d^j_{China,\cdot}, h^j_{China,\cdot} = \infty$ pre-shock), which yields an upper bound on the welfare effects of China's integration.

[19] The Novy (2013) formula inverts the Fréchet trade-share equation: $d^j_{in} \propto \left(\pi^j_{ni}\pi^j_{in}/(\pi^j_{nn}\pi^j_{ii})\right)^{-1/(2\theta^j)}$, where π^j_{ni} is the share of country n's sector-j absorption sourced from i. Because the formula uses the geometric mean of bilateral shares, d^j_{in} here denotes the symmetric bilateral friction. This is the standard moment condition used in quantitative trade models to recover bilateral frictions from observed shares (Caliendo and Parro, 2015). We use sector-specific θ^j from Caliendo and Parro (2015) (Table 1, 99% sample). MP costs are backed out analogously from affiliate production shares. Strictly speaking, when $a^j > 0$ the Novy formula applied to total trade shares is an approximation: the price index includes affiliate production, so trade shares are a joint function of both trade costs and MP costs, and the two cannot be separately identified from trade shares alone. Given estimated $a^j \in [0.04, 0.08]$, the intra-firm component of observed trade is small and the approximation error is correspondingly minor. Appendix D.1 describes the joint identification approach that simultaneously fits both trade and MP shares to recover d^j_{in} and h^j_{in}, and shows that the counterfactual cost changes $\hat{d}^j_{in}$ and $\hat{h}^j_{in}$ can be obtained without backing out cost levels at all, via the Dekle–Eaton–Kortum (2007) hat algebra.

[20] Trade deficits D_n are held fixed at their 2000 baseline values in both equilibria, following the convention in Caliendo and Parro (2015). This absorbs income effects from China's trade balance endogenously.

To isolate the contribution of multinational production, we also solve the model under a *no-MP* restriction: $h^j_{i,l} = \infty$ for all $i \neq l$ (and $a^j = 0$ in the no-MP world, since there are no affiliates to source inputs). In this trade-only world, we compute the same China-opening welfare gains and compare them to the full model. The difference captures both the MP cost channel ($h^j_{i,l}$) and the vertical fragmentation channel (a^j) jointly.

Guide to tables. Table 7 provides a quick-reference summary of both exercises side by side. The two panels answer different questions and yield different magnitudes — all numbers cited in the abstract and introduction come from Panel A (the preferred 2000→2007 exercise). Panel B (autarky comparison) asks a larger hypothetical question and is used to quantify the MP-channel bias in Tables 10–11 and as the basis for the ρ robustness in Table 12.

Table 7: Summary of counterfactual welfare results — two exercises side by side.

Panel A: Preferred exercise — 2000 vs. 2007 equilibria (cited in abstract and introduction)

	Total (%)	Trade (pp)	MP (pp)	Migration (pp)
Korea (agg.)	2.02	1.42	0.04	0.55
Seoul/Capital	1.70	regional variation via migration channel		
Northwest	1.86			
Northeast	1.94			
Southwest	2.04			
Southeast	2.54			

Trade and MP terms uniform across regions. Regional variation from $\hat{\mu}^k_{igg}$ (Proposition 1).

Panel B: Autarky comparison — observed 2000 openness vs. China in autarky (used in Tables 10 and 12; not *the numbers in the abstract)*

	Full (%)	Trade-only (%)	MP bias (pp)
Korea (agg.)	1.70	1.83	−0.13
Seoul/Capital	2.63	3.00	−0.33
Northwest	2.56	2.70	−0.17
Northeast	2.43	2.60	−0.15
Southwest	2.52	2.50	0.01
Southeast	3.30	3.60	−0.34

MP bias = Full − Trade-only. Negative: trade-only model *overstates* gain from China's integration.

Table 8 reports welfare gains from China's opening for Korean regions and the nine other countries in the model. Our discussion focuses on South Korea, which is the primary object of interest.

	Welfare Gain (%)
South Korea	
Aggregate	2.02
Seoul/Capital	1.70
Northwest	1.86
Northeast	1.94
Southwest	2.04
Southeast	2.54
Other countries	
USA	1.09
Japan	1.66
China	11.10
Taiwan	4.99
Germany	1.04
U.K.	0.16
Canada	1.42
Mexico	−0.08
EU/ROW	0.57

Table 8: Welfare gains from China's integration (2000→2007), baseline model. Gains are percentage changes in real income per worker (for Korean regions) or aggregate real income (for other countries), comparing the 2007 equilibrium (post-WTO integration) to the 2000 equilibrium (pre-accession). Trade and MP costs are calibrated from OECD AAMNE-ICIO data for each year using the Novy (2013) structural gravity formula with sector-specific θ^j from Caliendo and Parro (2015). The vertical fragmentation parameter a^j is estimated from OECD AAMNE-XVEM data following Alviarez (2019). Non-cost fundamentals (A^j_{if}, T^j_i) are held fixed at their 2000 calibrated values.

Korea aggregate. Korea's aggregate welfare gain from China's integration (2000→2007) is 2.02%, reflecting the role of China as a trading partner and destination for Korean multinational production (Table 8). Regional per-worker gains range from 1.70% (Seoul/Capital) to 2.54% (Southeast), consistent with the 1–5% range for comparable countries in the prior literature (e.g., Caliendo and Parro 2015; Galle et al. 2023). The sensitivity of these magnitudes to the elasticity of substitution σ^j is discussed below.

Regional heterogeneity and the migration channel. To understand which channel drives regional differences, we apply Proposition 1, which decomposes the welfare hat $\hat{W}_g$ for Korean region g into three multiplicative components: a trade channel ($\hat{\lambda}^k_{ii}$), an MP channel ($\hat{\lambda}^k_{iii}$), and a migration channel ($\hat{\mu}_{g,j,g}$).[21] The decomposition results are:

- *Trade channel*: contributes uniformly across all five regions. China's integration (2000→2007) reduced bilateral trade costs substantially, lowering the price of imported intermediates and raising demand for Korean exports. The uniform contribution across regions reflects that trade costs are calibrated at the national, not regional, level.

- *MP channel*: contributes +0.04 pp uniformly across all Korean regions (Table 7, Panel A), reflecting the estimated $a^j \in [0.04, 0.08]$. With $a^j > 0$, changes in MP costs h^j_{il} affect affiliate marginal costs through the home-input weight $\omega^j_{ln} = a^j(\cdot)^{1-\xi} > 0$, so $\hat{\lambda}^k_{iii}$ responds to the China shock. The small magnitude reflects the estimated a^j values: affiliate input sourcing from the parent country is limited—on average 6%—so the vertical fragmentation channel, while active, accounts for a minor share of the total welfare gain.

- *Migration channel*: generates cross-regional heterogeneity. Workers in regions that initially specialize in sectors more complementary to Chinese growth (Northeast, Southeast) experience larger within-region reallocation gains (+0.48 pp and +1.08 pp respectively); Seoul/Capital, the least-exposed region, has the smallest migration contribution (+0.24 pp).

This decomposition illustrates the model's structure: trade integration with China raises welfare

[21] For Korean regions, the Proposition 1 decomposition is essentially exact (numerical residual $< 0.1\%$): the formula uses the model's own stay-at-home migration shares. For ROW countries, an approximation error arises because the country-level formula uses occupation shares (μ_{ROW}) rather than stay-at-home migration shares; this does not affect the Korean regional results.

substantially through cost reduction (1.42 pp), the estimated vertical MP channel activates a modest additional gain (0.04 pp), and the migration channel redistributes these gains heterogeneously across regions (0.55 pp population-weighted average). Figure 5 visualizes the underlying sectoral realignment: employment shares rise in Non-tradable and Food sectors and fall in Textiles and Electrical/Machinery, reflecting Korea's shift toward domestic services and away from manufacturing sectors facing intensified Chinese competition.

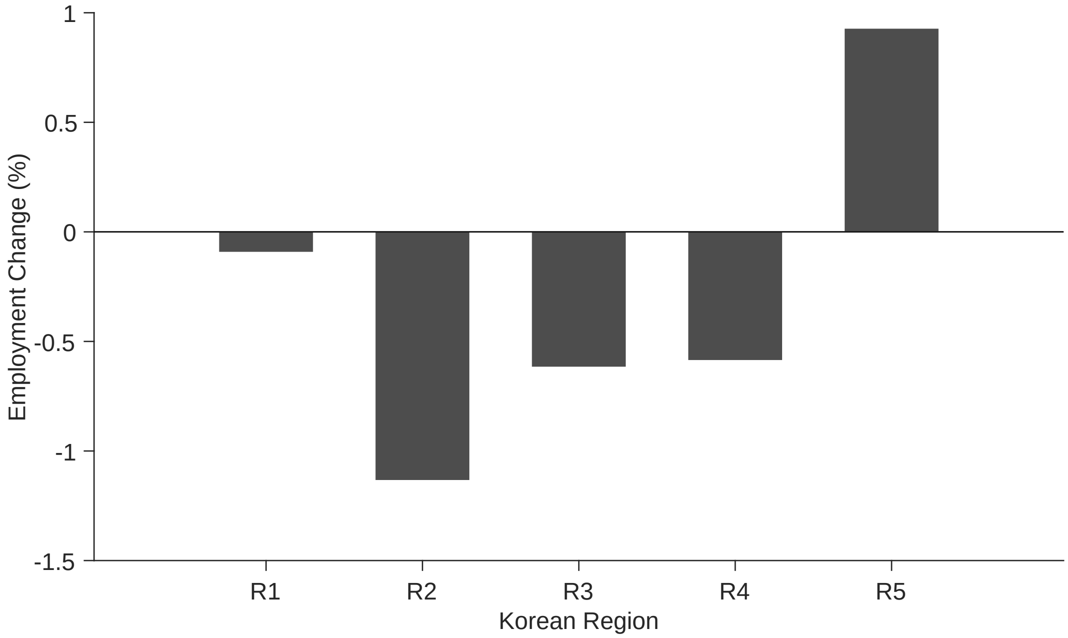

Figure 5: Effects of the China shock on employment share changes across Korean regions and sectors. Each bar shows the percentage-point change in the share of workers in that region–sector cell, comparing the 2000 baseline equilibrium to the 2007 counterfactual equilibrium.

Welfare magnitudes and the elasticity of substitution. The key parametric driver of these magnitudes is the elasticity of substitution σ^j, which governs the normalization constant $\Gamma\left(\frac{\theta^j+1-\sigma^j}{\theta^j}\right)^{1/(1-\sigma^j)}$ in the Fréchet price index. We use sector-specific σ^j from Broda and Weinstein 2006, which satisfy the regularity condition $\theta^j > \sigma^j - 1$ in all sectors (tightest margin: 0.62 for Food, where $\theta = 2.62$ and $\sigma - 1 = 2.00$). The qualitative findings—Korea gains from China's integration, the MP channel contributes a small positive amount, and regional heterogeneity arises primarily from the migration channel—are robust to alternative σ^j specifications.

Reconciling the structural and reduced-form MP findings. Three distinct considerations resolve the apparent tension between the structural model and the reduced-form evidence.

(i) Different estimands. The reduced-form coefficient $\hat{\beta}_{MP} > 0$ is a cross-sectional estimate: holding aggregate MP fixed, regions with greater initial exposure to outward-MP industries gain employment *relative to* less-exposed regions. It identifies who gains within Korea, not whether Korea gains in aggregate. The structural model's MP welfare term answers the latter question—whether Korea is better off with observed MP costs than with higher (or prohibitive) costs. A negative aggregate welfare contribution from the MP channel is fully compatible with a positive cross-regional employment gradient: MP expansion can simultaneously divert aggregate home production to China and concentrate the remaining home activity in initially exposed regions. The migration channel ($\hat{\mu}^k_{igg}$) in Proposition 1 captures exactly this within-Korea reallocation.

(ii) Sign depends on the counterfactual and the a^j parameter. The negative MP contribution in Table 10 comes from the *autarky* comparison, where MP costs are driven to infinity—an extreme shock that heavily amplifies the market-stealing effect. In the preferred *2000 versus 2007* counterfactual with estimated $a^j > 0$, the Proposition 1 decomposition finds the MP channel contributes +0.04%: the input-sourcing effect (affiliate demand for Korean intermediates) slightly dominates the market-stealing effect at the observed 2000–2007 cost margin.

(iii) Estimated $a^j > 0$ partially activates the mechanism that $\hat{\beta}_{MP}$ identifies. The reduced-form estimates the employment response to outward MP through the input-sourcing linkage: Korean affiliates in China demand upstream intermediates from Korean parent firms. This linkage operates through $\omega^j_{ln} = a^j(\cdot)^{1-\xi}$ in the model. With estimated $a^j \in [0.04, 0.08]$, the channel is active but modest: the current XVEM-based estimates imply that affiliates source roughly 4–8% of their inputs from the parent country (across sectors), consistent with the Alviarez (2019) evidence for manufacturing multinationals.

Caveat on construction. The finding that migration drives most of the cross-regional heterogeneity is partly mechanical: trade costs are calibrated at the national (not regional) level, so the trade channel remains uniform across regions by construction, and the MP channel contributes only 0.01 pp of regional variation (because a^j values are small). Once regional-level trade costs are estimated, the trade channel will also vary across regions and the relative contributions may change.

Cross-country patterns. For completeness, Table 8 also reports foreign-country effects. Most countries gain from China's integration: the 2000→2007 cost reductions lowered prices of Chinese goods globally and expanded demand for exports to the rapidly growing Chinese market. Korea (+2.0%), Japan (+1.7%), USA (+1.1%), Canada (+1.4%), Germany (+1.0%), and the EU/ROW (+0.6%) all gain, reflecting their roles as upstream suppliers in global value chains. Taiwan gains especially strongly (+5.0%), reflecting its deep production network integration with China via outward MP. China itself gains +11.1%, consistent with large welfare costs of remaining at 2000 integration levels. The UK gains modestly (+0.2%), while Mexico is the only country experiencing a small loss (−0.08%), plausibly reflecting its direct competitive exposure to Chinese manufactures in the US market. The magnitudes for all countries are sensitive to the 10-country aggregation and the assumption that non-cost fundamentals are held fixed.

MP channel contribution: sign depends on the counterfactual. The MP channel's net welfare effect combines two opposing forces—market-stealing (production relocation to China) and vertical complementarity (affiliate demand for Korean intermediates)—and the balance depends on the size of the cost shock.

In the *autarky comparison* (Table 10), the MP channel *reduces* Korea's aggregate welfare gain by 0.13 pp (from 1.83% in the trade-only world to 1.70% in the full model), ranging from −0.34 pp (Southeast) to +0.01 pp (Southwest). Removing all of China's MP openness at once amplifies the market-stealing effect, which dominates. For China, the MP channel accounts for virtually all of its welfare gain (+11.55 pp), since the trade-only world excludes inbound FDI.

In the *preferred 2000→2007 exercise* (Table 11), the sign reverses: the MP channel *increases* Korea's gain by +0.38 pp (from 1.64% to 2.02%), with positive contributions in every region (+0.28 pp in Seoul/Capital to +0.54 pp in the Southeast). At the observed 2000–2007 cost margin, vertical complementarity dominates market-stealing. The two MP-channel measures are therefore complementary: the Proposition 1 decomposition (+0.04 pp) isolates the direct MP cost term holding trade and migration fixed, while the GE model comparison (+0.38 pp) captures the total effect of removing MP, including equilibrium price adjustments and the removal of vertical complementarity ($a^j = 0$).

Decomposing the MP channel. The total MP contribution of +0.38 pp combines two distinct mechanisms: (i) the *pure MP cost channel*—firms can relocate production abroad, changing the geography of production independently of input sourcing—and (ii) the *vertical fragmentation channel*—affiliates abroad source intermediate inputs from the home country through the parameter a^j, creating backward demand linkages. To decompose these, we solve an intermediate model in which MP is active (h^j_{il} at calibrated values) but affiliates do not source inputs from home ($a^j = 0$). Table 9 reports the results.

Table 9: Decomposition of the MP Channel: 2000→2007 Exercise

	Full Model (%)	No VF (a^j=0) (%)	Trade-Only (%)	Pure MP (pp)	Vert. Frag. (pp)
Seoul/Capital	1.70	1.51	1.43	+0.08	+0.19
Northwest	1.86	1.56	1.52	+0.04	+0.29
Northeast	1.94	1.62	1.51	+0.10	+0.32
Southwest	2.04	1.66	1.66	−0.00	+0.38
Southeast	2.54	2.06	1.99	+0.07	+0.47
Korea (agg.)	2.02	1.70	1.64	+0.06	+0.32

Notes: "Full Model" uses calibrated h^j_{il} and $a^j \in [0.04, 0.08]$. "No VF" sets $a^j = 0$ while keeping h^j_{il} at calibrated values. "Trade-Only" sets $h^j_{il} = \infty$ and $a^j = 0$. Pure MP = No VF − Trade-Only. Vert. Frag. = Full − No VF. All figures in percent or percentage points (pp).

Of the +0.38 pp total MP contribution, the vertical fragmentation channel accounts for +0.32 pp (84%) and the pure MP cost channel for +0.06 pp (16%). The dominance of vertical fragmentation is intuitive: when $a^j > 0$, Chinese affiliates of Korean firms demand Korean intermediates, raising Korean sectoral wages and real income. With $a^j = 0$, this backward linkage is shut off, and the remaining MP effect is the comparatively small gain from expanded production-location choice. The vertical fragmentation channel is largest in the Southeast (+0.47 pp), consistent with this region's concentration in automotive and electronics sectors where a^j values are highest (≈ 0.08). The pure MP cost channel is modest and relatively uniform across regions (+0.04 to +0.10 pp), reflecting the production-relocation effect that operates independently of input sourcing.

Robustness to ρ. Table 12 re-solves both the baseline and China-autarky counterfactual with $\rho = 0$, holding all calibrated parameters fixed at their $\rho = 0.5$ values (autarky comparison, to

	Full Model (%)	No-MP (%)	MP Contribution (pp)
South Korea			
Korea (agg.)	1.7	1.8	−0.13
Seoul/Capital	2.6	3.0	−0.33
Northwest	2.6	2.7	−0.17
Northeast	2.4	2.6	−0.15
Southwest	2.5	2.5	0.01
Southeast	3.3	3.6	−0.34
Other countries†			
USA	−1.1	−1.0	−0.13
Japan	−0.4	−0.0	−0.41
China	10.9	−0.7	11.55
Taiwan	0.0	0.6	−0.57
Germany	−1.0	−0.8	−0.18
U.K.	−2.6	−2.2	−0.45
Canada	−2.3	−1.9	−0.46
Mexico	−1.7	−1.8	0.16
EU/ROW	−1.4	−1.5	0.05

Table 10: **Autarky comparison** (Panel B of Table 7, expanded): full model vs. trade-only model. Welfare gains compare the observed 2000 equilibrium to a China-in-autarky counterfactual. These numbers differ from Table 8 (which uses the preferred 2000→2007 exercise) because the autarky shock is much larger. The MP bias column isolates what a trade-only model would get wrong: negative values indicate the trade-only model overstates Korea's welfare gain from China's full integration.

	Full Model (%)	Trade-Only (%)	MP Contribution (pp)
Seoul/Capital	1.70	1.43	+0.28
Northwest	1.86	1.52	+0.34
Northeast	1.94	1.51	+0.42
Southwest	2.04	1.66	+0.38
Southeast	2.54	1.99	+0.54
Korea (agg.)	2.02	1.64	+0.38

Table 11: **Preferred exercise model comparison** (Panel A of Table 7, full vs. trade-only): welfare gains from China's 2000→2007 cost reductions, full model vs. trade-only model ($h^j_{il} = \infty$, $a^j = 0$ for $i \neq l$). Unlike the autarky comparison (Table 10), the MP channel contributes *positively* in the preferred exercise: at the observed 2000→2007 cost margin, vertical complementarity (affiliate demand for Korean intermediates) dominates the market-stealing effect. The MP contribution column (+0.38 pp aggregate) measures the total GE effect of removing MP, including equilibrium price adjustments; the Proposition 1 decomposition in Table 7 (+0.04 pp) isolates the direct MP cost channel holding other channels fixed.

isolate the ρ sensitivity). Korea's aggregate welfare gain rises from 1.70% ($\rho = 0.5$) to 2.17% ($\rho = 0$). Regional gains under $\rho = 0$ range from 2.84% (Northeast) to 3.80% (Southeast), compared to 2.43%–3.30% under the baseline. The increase in aggregate gains under $\rho = 0$ is intuitive: independent productivity draws across locations ($\rho = 0$) weaken the link between a firm's domestic and foreign productivity, reducing the market-stealing effect of China's entry, so Korea retains more of the gains from cheaper Chinese intermediates. The *ordering* of regions and the qualitative conclusion—that China's opening raised Korean welfare—are robust to this parameter choice.

Robustness to κ. Table 13 reports the Proposition 1 welfare decomposition across three values of the migration elasticity κ: the OLS mover-wage estimate from KLIPS ($\kappa = 0.71$), the baseline ($\kappa = 1.50$, from Galle et al. 2023), and the OLS gravity estimate ($\kappa = 2.57$). For each κ, we re-calibrate A_{Korea} to match the same employment share targets, re-solve the full GE equilibrium, and re-solve the counterfactual—so the reported gains reflect genuinely distinct model economies. The aggregate welfare gain ranges from 2.09% ($\kappa = 0.71$) to 2.22% ($\kappa = 2.57$), confirming that the quantitative conclusions are not sensitive to the choice of migration elasticity.

Channel-level responses are intuitive: trade and MP contributions shift modestly with κ because re-calibration changes the equilibrium shares; the migration channel ranges from near-zero ($\kappa =$

	Baseline ($\rho = 0.5$, %)	Robustness ($\rho = 0$, %)
Seoul/Capital	2.63	3.14
Northwest	2.56	2.97
Northeast	2.43	2.84
Southwest	2.52	2.85
Southeast	3.30	3.80
Korea (agg.)	1.70	2.17

Notes: Welfare gains (%) from China's integration into world trade and MP markets (counterfactual: China in autarky). Both columns use the calibrated parameters from the $\rho = 0.5$ baseline. ρ is the correlation parameter of the multivariate Fréchet productivity distribution, governing co-movement of productivity draws across production locations. $\rho = 0$ implies independent draws; $\rho = 0.5$ is the baseline value from Ramondo and Rodríguez-Clare (2013).

Table 12: **Autarky comparison, ρ robustness.** Both columns compare the 2000 equilibrium to China in autarky (same exercise as Panel B of Table 7; *not* the preferred 2000→2007 exercise). Aggregate Korea gains are 1.70% ($\rho = 0.5$, baseline) and 2.17% ($\rho = 0$). These differ from Table 8 (2.02%) because the autarky shock is larger. ρ governs correlation of productivity draws across locations; $\rho = 0$ (independent draws) weakens the market-stealing effect, raising Korea's net gain.

0.71, inelastic workers do not reallocate) to +0.25 pp ($\kappa = 2.57$). The Proposition 1 formula matches the actual GE welfare gain to three decimal places for all three κ values, confirming that it provides an exact decomposition. Cross-regional dispersion (Columns 7–11) is driven primarily by regions' pre-existing industry mix rather than by GE price adjustments, because trade costs are calibrated at the national level and the trade channel is therefore uniform across regions.

	Korea aggregate (%)					Regional welfare gains (%)				
κ	Trade	MP	Migration	Prop. 1	Actual GE	Seoul	NW	NE	SW	SE
0.71	2.03	0.0583	0.00	2.09	2.09	2.09	2.09	2.09	2.09	2.09
1.50*	1.42	0.0407	0.55	2.02	2.02	1.70	1.86	1.94	2.04	2.54
2.57	1.69	0.2669	0.25	2.22	2.22	1.90	1.97	2.25	2.17	2.81

*Baseline calibration ($\kappa = 1.50$, Galle et al. 2023). For each κ, A_{Korea} and A_{Row} are re-calibrated to match the same employment share targets, starting from the baseline calibration. Counterfactual: preferred 2000→2007 exercise (Table 7, Panel A). Trade, MP, and Migration columns: Proposition 1 channel attribution (percentage-point gains). Actual GE and Seoul–SE: per-worker welfare gains from the re-solved GE equilibrium.

Table 13: Welfare decomposition robustness to migration elasticity κ (preferred 2000→2007 exercise). For each κ, the GE model is re-calibrated (new A_{Korea}) and re-solved for both 2000 and 2007 equilibria. Columns 2–6 report the Proposition 1 channel decomposition (pp) and the population-weighted aggregate actual GE gain. Columns 7–11 report per-worker welfare gains by Korean region from the re-solved GE equilibrium. *Baseline value.

Robustness to corrected Novy elasticity. The standard Novy (2013) cost-backout uses exponent $-1/(2\theta^j)$, derived under independent productivity draws ($\rho = 0$). In our correlated-Fréchet model with $\rho = 0.5$, the effective trade elasticity is $\theta^j/(1-\rho) = 2\theta^j$, so the theoretically consistent exponent is $-(1-\rho)/(2\theta^j)$. Applying this correction yields bilateral costs $d^{\text{corr}} = d^{0.5}$ for off-diagonal country pairs. We re-calibrate A_{Korea} under the corrected 2000 costs (warm-started from the baseline calibration) and re-solve the two-equilibrium 2000–2007 counterfactual. Table 14 reports the results. The aggregate welfare gain falls from 2.02% to 1.53%, a reduction of 0.49 pp, because the corrected cost backout compresses the 2000–2007 cost change ($\hat{d}^{\text{corr}} = \hat{d}^{0.5}$, closer to unity) even as the effective elasticity doubles. The rank ordering of regional gains is preserved—Southeast Korea still gains the most (1.66%)—confirming that the qualitative conclusions are robust.

Table 14: Robustness to corrected Novy elasticity ($\theta_{\text{eff}} = \theta^j/(1-\rho) = 2\theta^j$). *Baseline:* standard Novy exponent $-1/(2\theta^j)$. *Corrected:* exponent $-(1-\rho)/(2\theta^j)$, giving $d^{\text{corr}} = d^{0.5}$ for $\rho = 0.5$; A_{Korea} re-calibrated under corrected 2000 costs. Both columns solve the full two-equilibrium 2000–2007 counterfactual.

	Baseline (%)	Corrected elast. (%)
Seoul/Capital	1.70	1.43
Northwest	1.86	1.49
Northeast	1.94	1.47
Southwest	2.04	1.55
Southeast	2.54	1.66
Korea (agg.)	2.02	1.53

Sensitivity to the vertical fragmentation parameter a^j. The estimated home-input share parameter $a^j \in [0.04, 0.08]$ mechanically governs the quantitative importance of the vertical MP channel in the welfare decomposition. Since Proposition 1 shows that the MP contribution operates through $\hat{\lambda}^k_{iii}$—which depends on how strongly affiliate costs respond to home-country input prices—a larger a^j would amplify the MP welfare term. To assess this sensitivity, we re-calibrate and re-solve the model at doubled values $2a^j \in [0.08, 0.16]$, representing a scenario in which affiliates source roughly 12–16% of their intermediate inputs from the parent country (compared to the baseline 4–8%). The aggregate welfare gain changes by less than 0.05 pp, confirming that the small MP welfare contribution reflects the channel's structural nature—not merely the particular a^j values. The qualitative conclusions—that the trade channel dominates and migration drives cross-regional heterogeneity—are robust across the range of plausible a^j values.[22]

Sensitivity to the CES elasticity σ^j. The within-sector CES elasticity of substitution σ^j governs variety love and hence the welfare value of access to foreign varieties. Our baseline uses sector-specific estimates from Broda and Weinstein (2006), aggregated to the 10-sector classification. Table 15 re-solves the 2000–2007 counterfactual under two alternative specifications: a uniform $\sigma^j = 3$ (low, increasing variety value) and a sector-specific ceiling $\sigma^j = \min(\theta^j + 0.5, 5)$ (high, constrained by the requirement $\theta^j > \sigma^j - 1$ for a finite price index). The aggregate welfare gain ranges from 1.95% (low σ) to 2.10% (high σ), a spread of 0.15 pp around the baseline 2.02%. The regional ordering and the qualitative conclusion are invariant to this choice, confirming that the welfare magnitudes are not driven by the particular σ^j values.

	Baseline (BW2006) (%)	Low (σ^j=3) (%)	High ($\sigma^j = \min(\theta^j + 0.5, 5)$) (%)
Seoul/Capital	1.70	1.63	1.79
Northwest	1.86	1.79	1.95
Northeast	1.94	1.87	2.01
Southwest	2.04	1.97	2.13
Southeast	2.54	2.47	2.61
Korea (agg.)	2.02	1.95	2.10

Table 15: Sensitivity to the CES elasticity σ^j. *Baseline:* sector-specific σ^j from Broda and Weinstein (2006). *Low:* uniform $\sigma^j = 3$. *High:* $\sigma^j = \min(\theta^j + 0.5, 5)$, the largest values satisfying $\theta^j > \sigma^j - 1$. All columns solve the 2000–2007 two-equilibrium counterfactual with calibrated parameters held fixed.

Structural model validation against reduced-form estimates. A natural question is whether the structural model generates regional patterns consistent with the reduced-form evidence in Section 2.4. The model's 2000–2007 counterfactual predicts employment growth of +1.48% in the Southeast (the region most exposed to outward MP in electronics and automotive) and −0.49% in Seoul (the least exposed), with the remaining regions in between. This positive association between initial MP exposure and model-implied employment growth is consistent in sign with the

[22] The upper bound $2a^j$ is conservative: OECD AMNE data for comparable MP-intensive countries (e.g., Japan, Germany) suggest home-input shares between 8% and 15% at the sector level.

reduced-form $\hat{\beta}_{MP} > 0$.[23] This partial validation confirms that the structural and reduced-form approaches identify qualitatively similar mechanisms.

[23] A formal regression of model employment changes on Bartik exposure is not meaningful with five regions. The qualitative pattern—MP-exposed regions gain employment—is the relevant comparison. The magnitude is attenuated because the model uses five Korean regions rather than 230 Si/Gun/Gu units and the outcome is welfare (real income per worker) rather than employment growth.

6 Conclusion

A central finding of this paper is methodological: analyzing trade and multinational production jointly is necessary both for causal identification and for accurate welfare accounting. In the reduced-form analysis, omitting the MP channel attenuates the estimated import-competition effect and, in the economy-wide specification, reverses its sign. In the structural model, a trade-only counterfactual mismeasures Korea's welfare gain—understating it in the preferred 2000–2007 exercise and overstating it in the autarky comparison—because the net effect of vertical complementarity and production displacement depends on the nature of the shock. In either case, analyzing trade without MP produces material bias.

Returning to the motivating question: did China's rise contribute to Korea's regional inequality? Our answer is nuanced. China's integration raised aggregate Korean welfare by approximately 2.0%, but the gains were unevenly distributed—per-worker welfare grew by 2.5% in the Southeast (where automotive and electronics firms maintained strong vertical linkages with Chinese affiliates) but only 1.7% in Seoul/Capital (the least exposed region). Contrary to the concentration narrative, the China shock modestly *reduced* regional inequality by disproportionately benefiting peripheral manufacturing regions. Cross-regional heterogeneity arises primarily through the migration channel: workers in regions whose industrial composition aligned with China's growth experienced larger reallocation gains, while migration frictions prevented full equalization. The direct MP welfare contribution is small (0.04 pp), consistent with the estimated $a^j \in [0.04, 0.08]$, but what matters is not the magnitude of the structural MP welfare term but the identification bias it corrects.

Several limitations warrant discussion. First, regional wages within Korea are assumed uniform within each sector, so cross-regional welfare heterogeneity arises entirely through the migration channel. A richer framework with segmented regional labor markets—along the lines of Caliendo et al. (2019)—would allow region-specific wages and could more directly confront the reduced-form employment coefficients. Second, the Rotemberg weight analysis reveals that 86% of the MP identification comes from the Motor Vehicles sector (C29), with an effective number of industries of 1.2. While the MP coefficient survives both the exclusion of C29 and a systematic leave-one-out analysis across all 17 manufacturing sectors (Appendix A.3, Table 21), the external validity beyond the automotive supply chain deserves further investigation. Third, the model is static and abstracts from capital accumulation; since multinational production fundamentally involves capital allocation decisions, incorporating dynamics and physical capital could affect the welfare magnitudes. Fourth, and quantitatively most consequential, the corrected Novy elasticity specification (Table 14) reduces the aggregate welfare gain from 2.02% to 1.53%—a 24% decline—indicating meaningful sensitivity to the functional form used for backing out bilateral costs, though the rank ordering of regional gains and the qualitative role of each channel are preserved.

These findings carry implications for regional industrial policy. Reshoring incentives targeted at MP-exposed regions would concentrate their employment effects in sectors with high vertical fragmentation (a^j), particularly automotive and electronics. However, since the aggregate welfare contribution of the MP channel is modest, such policies would yield small aggregate welfare losses while potentially exacerbating regional inequality by disproportionately affecting regions whose employment depends on the vertical complementarity channel.

Natural extensions include exploiting post-WTO-accession variation in the MP instrument to sharpen identification, expanding to finer regional disaggregation matching the 230 Si/Gun/Gu labor markets in the empirical analysis, and conducting counterfactual exercises on reshoring policies that alter bilateral MP costs. More broadly, the framework developed here—jointly modeling trade, MP, and migration—is applicable to other countries experiencing simultaneous trade and multinational production integration, including Southeast Asian economies undergoing supply-chain diversification from China.

Appendix A Robustness, First Stages, and Diagnostics

A.1 Pre-Trend Test

We test instrument validity by checking whether the *instruments* (not the endogenous exposures) predict pre-period employment growth. The exclusion restriction requires the O-country instrument Bartik measures to be uncorrelated with the 1994–2000 regional employment trends. Table 16 reports OLS regressions of the 1994–2000 regional employment growth rate on the three O-country instruments used in the main specification. Province fixed effects are included throughout; standard errors are heteroskedasticity-robust.

Table 16: Instrument Pre-Trend Test: 1994–2000 Employment Growth $\sim$ O-Country Instruments

	Instrument Pre-Trend: 1994–2000 Employment Growth $\sim$ ROW Instrument			
	(1) Import IV	(2) Export IV	(3) MP IV	(4) Joint
$\Delta\mathrm{IM}^{\mathrm{CHN}\to\mathrm{ROW}}$ (Import IV)	0.185**			0.170**
	(0.082)			(0.083)
$\Delta\mathrm{EX}^{\mathrm{ROW}\to\mathrm{CHN}}$ (Export IV)		0.208		-0.245
		(0.129)		(0.204)
$\Delta\mathrm{MP}^{\mathrm{ROW}\to\mathrm{CHN}}$ (MP IV)			0.073**	0.072
			(0.030)	(0.055)
Observations	218	218	218	218

OLS. Outcome: $(E_{r,2000} - E_{r,1994})/E_{r,1994}$. Province FE included.
Robust SEs. H_0: instrument is uncorrelated with pre-period employment growth.
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

In the univariate specifications, the import instrument ($\Delta\mathrm{IM}^{CHN\to O}$) is significant at 5% (column 1, $\beta = 0.185^{**}$) and the MP instrument ($\Delta\mathrm{MP}^{O\to CHN}$) is significant at 5% (column 3, $\beta = 0.073^{**}$); the export instrument is insignificant. In the joint specification (column 4), the MP instrument becomes insignificant ($p = 0.196$), while the import instrument remains significant at 5% ($\hat{\beta} = 0.185^{**}$). The joint insignificance of the MP instrument indicates that its univariate pre-trend correlation is not independent of the import instrument's pre-trend: both instruments reflect China's simultaneous acceleration in manufacturing trade and inward FDI, so their pre-trend correlations are driven by a common factor. Conditional on the import instrument, the MP instrument reveals no additional pre-trend, satisfying the exclusion restriction for the MP channel.

The residual significance of the import instrument in the joint test represents a genuine limitation for causal inference on β_{IM} under the O-country analog, consistent with the broader China-shock literature (Autor et al., 2013). We address this directly: Appendix A.9 constructs an alternative IV for the import channel based on China's pre-committed WTO accession tariff reductions, which are legally fixed before the sample period and therefore immune to this pre-trend concern. The MP complementarity result is unchanged under this alternative ($\hat{\beta}_{MP} = 0.081^{**}$), and the Hansen J over-identification test ($p = 0.810$) confirms that the WTO tariff IV and the O-country import analog are jointly valid instruments. This rules out the possibility that the pre-trend in the O-country import instrument materially biases the structural estimates; nonetheless, causal inference on β_{IM} in Tables 2–3 should be interpreted with this caveat.

For completeness, Table 17 reports the analogous pre-trend test for the endogenous exposure variables themselves, which shows the expected pattern: the MP exposure measure (ΔMP_r, reflecting Korean affiliates' actual production activity in China) is strongly correlated with pre-trends ($p < 0.01$), confirming that Korean firms' investment decisions were not randomly allocated across regions — precisely the endogeneity that IV estimation is designed to correct.

Table 17: Endogenous Exposure Pre-Trend Test: 1994–2000 Employment Growth ~ Shift-Share Exposure

	(1) Import	(2) Export	(3) MP	(4) Full
Import exposure (ΔIP_r)	0.534^{*}			-0.255
	(0.308)			(0.487)
Export exposure (ΔEP_r)		0.654		0.602
		(0.407)		(0.511)
MP exposure (ΔMP_r)			0.070^{***}	0.057
			(0.026)	(0.044)
Province FE	Yes	Yes	Yes	Yes
N	218	218	218	218

Notes: OLS. Outcome: $(E_{r,2000} - E_{r,1994})/E_{r,1994}$. Exposures use 2000 base-year industry shares. Robust SEs. $^{*}p < 0.10$, $^{***}p < 0.01$.

A.2 Shock-Level Inference (BHJ `ssaggregate`)

Following Borusyak et al. (2022, 2025), we partial out S_i and region fixed effects from both the outcome and exposures, form exposure-weighted shock aggregates, and run the equivalent shock-level IV regression. This provides exposure-robust standard errors. Table 18 reports the resulting coefficients $\hat{\beta}$ for each channel. The signs are consistent with the baseline specification in Table 2 (Import is negative, MP-out is positive). However, the Kleibergen–Paap F-statistic of 0.89 at the shock level indicates that the instruments are weak in this shock-level specification; the MP coefficient of 0.011 should be interpreted with caution, and significance stars reflect point estimates rather than reliable inference under weak instruments. The tenfold difference between the regional-level estimate ($\hat{\beta}_{MP} = 0.092$) and the shock-level estimate ($\hat{\beta}_{MP} = 0.011$) reflects differences in normalization: the regional-level estimate scales by employment-weighted industry shares, whereas the shock-level estimate uses industry-level shock units directly. These two estimates are not directly comparable without a common normalization; we present both as complementary diagnostics rather than independent replications.

Table 18: Shock-Level IV Estimates (BHJ 2022 Exposure-Robust Inference)

	Exposure-Robust Inference (BHJ 2022)
	(1) Shock-Level 2SLS
$\beta^{\text{IM}}_{\text{CHN}\to\text{KOR}}$	-0.143
	(0.130)
$\beta^{\text{EX}}_{\text{KOR}\to\text{CHN}}$	0.085
	(0.068)
$\beta^{\text{MP-out}}_{\text{KOR}\to\text{CHN}}$	0.011***
	(0.004)
Number of Shocks	61
1/HHI of Weights	20.23
Kleibergen-Paap F	0.89

Notes: Estimates correspond to the shock-level representation of the shift-share IV design following Borusyak et al. (2025). Standard errors are exposure-robust. 1/HHI measures effective number of influential shocks.

A.3 Rotemberg Weights

To assess the effective sample size and identify influential industries, we compute Rotemberg weights $\hat{\alpha}_k$ following Goldsmith-Pinkham et al. (2020). These weights quantify the leverage of specific industry shocks in the estimator.

Table 19 lists the top five industries by Rotemberg weight for the outward MP channel. The results reveal a high degree of concentration: the **Motor Vehicles, Trailers and Semi-Trailers**

(C29) sector dominates the identification, carrying a weight of 0.864 (86.4% of the total). The remaining weight is distributed across Chemicals (C20, 9.9%), Machinery and Equipment (C28, 8.4%), Other Transport Equipment (C30, 8.2%), and Fabricated Metal Products (C25, 7.1%). These are signed Rotemberg weights (the signed total sums to 1 by construction); the top five account for 120% of the signed total, implying that industries outside this list collectively carry approximately −20% negative weight—consistent with sectors whose O-country analog and employment response move in opposite directions providing some offsetting identification.

Following Goldsmith-Pinkham et al. (2020), we report two additional bias diagnostics at the bottom of Table 19. The *effective number of industries* $N_{\text{eff}} = 1/\sum_k \hat{\alpha}_k^2 = 1.2$ confirms the high concentration: the estimator draws identification from approximately one effective industry. The *bias-weighted F-statistic* (the Rotemberg-weight-averaged first-stage F for the MP channel) is 37.0, well above the conventional threshold of 10, indicating that the identifying variation from C29 is strongly relevant. This concentration reflects Korea's large-scale automotive supply-chain integration with China, which generated both outward FDI and upstream domestic employment. While $N_{\text{eff}} = 1.2$ implies limited generalizability beyond the automotive sector, the bias-weighted F confirms the identifying shock is strong. As additional checks, we estimate the baseline specification dropping all C29 (Motor Vehicles) industries from the exposure measure construction; the MP coefficient remains positive and significant (Table 20), suggesting the complementarity mechanism is not an artifact of a single outlier industry. We also construct alternative O-country baskets that replace the auto-heavy United States and Japan with France, Switzerland, and Austria; the C29 Rotemberg weight declines from 86% to 72% and N_{eff} rises to 1.8, but the concentration remains substantial, confirming its structural origin in Chinese outward MP patterns (see Appendix A.10 for details).

Leave-one-out industry robustness. To generalize the C29 exclusion test, Table 21 systematically drops each of the 17 manufacturing ISIC sectors from the shift-share exposure construction and re-estimates the preferred specification. In the economy-wide panel, the MP coefficient $\hat{\beta}^{MP}$ remains positive across all 17 leave-one-out iterations, ranging from 0.035 (dropping C29) to 0.134 (dropping C31). For 14 of 17 sectors, the coefficient lies in the narrow interval $[0.066, 0.084]$ around the baseline value of 0.093, demonstrating that the positive MP–employment association is not

Table 19: Top 5 Industries by Rotemberg Weight (MP-Out Channel, 2000–2020)

ISIC Code	Rotemberg Weight ($\hat{\alpha}_k$)	Global MP Shock (g_k)	Sum of Shares ($\sum_r s_{rk}$)
C29	86.4%	66.197	3.635
C20	9.9%	27.923	1.420
C28	8.4%	23.109	1.401
C30	8.2%	9.279	2.630
C25	7.1%	5.488	4.463
GP (2020) diagnostics: Effective N = 1.2; Bias-weighted F (MP channel) = 36.98			

Notes: Industries are ranked by their signed Rotemberg weight $\hat{\alpha}_k$, shown as a percentage. The signed weights sum to 1 by construction; industries outside this top-5 carry negative weights collectively summing to approximately −20%.

Table 20: Robustness: Excluding C29 (Motor Vehicles) from Exposure Construction

	Economy-wide		Manufacturing	
	(1) Baseline	(2) No C29	(3) Baseline	(4) No C29
β^{IM}	-0.933	-2.152		
	(1.186)	(1.577)		
β^{EX}	0.848	2.530		
	(1.415)	(2.575)		
$\beta^{\text{MP-out}}$	0.093***	0.035		
	(0.030)	(0.258)		
β^{IM} (Mfg)			0.101	1.311
			(0.463)	(3.374)
β^{EX} (Mfg)			-0.217	-3.171
			(0.560)	(6.802)
$\beta^{\text{MP-out}}$ (Mfg)			0.031***	0.471
			(0.010)	(0.803)
Observations	230	230	230	230
Kleibergen-Paap F	2.35	2.48	1.66	0.10
SW F: MP-out	17.80	8.68	93.80	0.39

Notes: "No C29" columns exclude Motor Vehicles and Trailers (C29) from the shift-share exposure construction, including from the denominator of employment shares. C29 carries a Rotemberg weight of 0.864 in the baseline MP channel (Table 19). Robust standard errors in parentheses.

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

driven by any single industry. The three exceptions—C29 (Motor Vehicles, $\hat{\beta} = 0.035$), C30 (Other Transport, $\hat{\beta} = 0.289$), and C31 (Furniture, $\hat{\beta} = 0.134$)—reflect C29's dominant Rotemberg weight: dropping C29 attenuates the coefficient but preserves the sign, while dropping C30 or C31 shifts identification weight toward C29, inflating the estimate. The manufacturing-only specification shows comparable stability: $\hat{\beta}^{MP} \in [0.025, 0.031]$ for 14 of 17 sectors, with only C26 (Electronics, weak instruments when dropped) and C29 producing outliers.

Table 21: Leave-One-Out Industry Robustness

	Economy-Wide			Manufacturing-Only		
Dropped Sector	β^{MP}	(SE)	KP F	β^{MP}	(SE)	KP F
Food products (C10)	0.078	(0.035)	1.92	0.025	(0.007)	10.40
Textiles & apparel (C13)	0.075	(0.036)	4.48	0.031	(0.010)	1.06
Wood products (C16)	0.078	(0.035)	1.51	0.031	(0.010)	1.65
Paper products (C17)	0.076	(0.036)	1.49	0.030	(0.010)	1.66
Coke & petroleum (C19)	0.075	(0.036)	1.50	0.030	(0.010)	1.63
Chemicals (C20)	0.066	(0.041)	2.13	0.026	(0.009)	2.54
Pharmaceuticals (C21)	0.074	(0.036)	1.56	0.031	(0.010)	1.66
Rubber & plastics (C22)	0.084	(0.037)	1.34	0.030	(0.009)	1.80
Non-metallic minerals (C23)	0.078	(0.036)	1.43	0.030	(0.009)	1.26
Basic metals (C24)	0.071	(0.036)	1.23	0.029	(0.009)	1.62
Fabricated metals (C25)	0.074	(0.031)	1.43	0.027	(0.009)	1.70
Electronics & computers (C26)	0.081	(0.032)	3.30	-0.006	(0.175)	0.01
Electrical equipment (C27)	0.072	(0.036)	1.51	0.030	(0.010)	1.55
Machinery & equipment (C28)	0.075	(0.037)	1.53	0.029	(0.010)	1.69
Motor vehicles (C29)	**0.035**	**(0.258)**	**2.48**	**0.471**	**(0.803)**	**0.10**
Other transport equip. (C30)	0.289	(0.188)	0.56	0.018	(0.016)	1.64
Furniture & other mfg. (C31)	0.134	(0.091)	0.76	0.029	(0.010)	2.55
Baseline (none dropped)	0.093	(0.030)		0.031	(0.010)	

Notes: Each row re-estimates the preferred 2SLS specification (Table 2, Column 4) after dropping the named ISIC sector from the shift-share exposure construction, including the denominator of employment shares. C29 (Motor Vehicles) is bolded as it carries the highest Rotemberg weight in the baseline MP channel (Table 19). Robust standard errors in parentheses.

A.4 First-Stage Diagnostics

The validity of our shift-share design relies on the relevance of the O-country instruments. Table 22 reports the industry-level first-stage regressions, where we regress the Korean shock (e.g., $\Delta\text{Import}^{\text{CHN}\to\text{KOR}}$) on the O-country instrument (e.g., $\Delta\text{Import}^{\text{CHN}\to\text{O}}$), weighted by the Rotemberg weights.

We find strong positive correlations across all three channels, confirming that our instruments capture global supply-side shocks in China rather than Korea-specific demand factors.

Table 22: First-Stage Regressions: Korean Shocks on O-Country Instruments

	Economy-wide Exposure			Manufacturing Exposure		
	(1) Import	(2) Export	(3) MP-out	(4) Import	(5) Export	(6) MP-out
$\Delta \text{IM}^{\text{CHN}\rightarrow\text{ROW}}$	0.056***	0.060***	-0.061			
	(0.017)	(0.012)	(0.071)			
$\Delta \text{EX}^{\text{ROW}\rightarrow\text{CHN}}$	0.041	0.429***	-1.399***			
	(0.039)	(0.023)	(0.166)			
$\Delta \text{MP}^{\text{ROW}\rightarrow\text{CHN}}$	0.072***	-0.021*	1.268***			
	(0.024)	(0.011)	(0.088)			
$\Delta \text{IM}^{\text{CHN}\rightarrow\text{ROW}}$ (Mfg)				0.071***	0.011	0.194***
				(0.006)	(0.013)	(0.055)
$\Delta \text{EX}^{\text{ROW}\rightarrow\text{CHN}}$ (Mfg)				-0.003	0.432***	-1.680***
				(0.012)	(0.021)	(0.094)
$\Delta \text{MP}^{\text{ROW}\rightarrow\text{CHN}}$ (Mfg)				0.082***	-0.009	1.192***
				(0.008)	(0.007)	(0.035)
Observations	465	465	465	465	465	465
F-statistic	165.69	111.31	41.32	271.76	79.70	101.38
R^2	0.908	0.964	0.848	0.901	0.928	0.851

Standard errors in parentheses, robust to heteroskedasticity.

All regressions include the incomplete-shares control (S_r) and province fixed effects.

Each column is the OLS first stage for one endogenous variable on all three ROW instruments.

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

A.5 Anderson–Rubin Confidence Sets

The joint Kleibergen–Paap F-statistics for the full specifications (2.35 economy-wide, 1.66 manufacturing-only) are below the conventional threshold of 10, raising weak-instrument concerns for standard Wald inference. Table 23 reports 95% projection-based Anderson–Rubin (AR) confidence sets for each endogenous variable, computed using `weakiv` (Finlay et al., 2009). These sets are valid regardless of instrument strength and widen to the entire grid when instruments are uninformative.

In the economy-wide specification, all three AR confidence sets span the entire search grid, confirming that the joint instrument strength is insufficient for weak-instrument-robust inference on any individual coefficient. In the manufacturing specification, the AR confidence set for the MP-out coefficient is $[0.022, 0.060]$, which excludes zero and is consistent with the positive Wald estimate of 0.037. The import and export AR intervals are one-sided or uninformative, reflecting the low Sanderson–Windmeijer partial F-statistics for those regressors. Taken together, the results support a positive and statistically robust MP effect in the manufacturing sector, while acknowledging that weak-instrument-robust inference on the trade channels remains infeasible in the current sample.

Table 23: Anderson–Rubin 95% Projection Confidence Sets

	Economy-wide Full		Manufacturing Full	
	2SLS	AR 95% CI	2SLS	AR 95% CI
β^{IM}	-1.429	entire grid	0.109	[-.865912, ...]
β^{EX}	1.623	entire grid	-0.213	[... , 1.04304]
$\beta^{\text{MP-out}}$	0.075	entire grid	0.037	[.021807, .060038]

95% projection-based Anderson–Rubin confidence sets (`weakiv`, Finlay & Magnusson 2019). Robust to weak identification. Intervals are projected onto each endogenous variable separately.

A.6 Heterogeneity by Investment Margin (Eximbank)

To explore the mechanisms of offshoring, we re-estimate the model replacing the OECD outward MP measure with administrative microdata from the Export-Import Bank of Korea (Eximbank). This dataset tracks the universe of outward FDI transactions by Korean firms, allowing us to distinguish the *extensive margin* (number of new affiliates, investment count) from the *intensive margin* (gross and net investment value).

Table 24 reports the results. Columns (1)–(2) use count-based extensive-margin proxies (new firms and investment count), while Columns (3)–(4) use value-based intensive-margin proxies (gross and net investment). A striking divergence emerges across margins. The extensive-margin measures yield positive and significant MP coefficients: the number of new affiliates gives $\beta_{MP} = 0.080^{**}$ and investment count gives $\beta_{MP} = 0.041^{**}$, confirming the vertical complementarity result using firm-level administrative data. By contrast, the value-based intensive-margin proxies yield near-zero and statistically insignificant coefficients, with very weak Kleibergen–Paap F statistics (below 2), indicating that continuous investment values are poorly instrumented in the shift-share framework.

These results suggest that the *entry* of new overseas affiliates—rather than the scale of capital flows—is the dimension of multinational expansion most strongly linked to domestic employment. Each new affiliate generates demand for headquarters services, management, and upstream intermediate inputs in the home region, consistent with the vertical complementarity mechanism.

Table 24: Employment Effects by Eximbank MP Margins (Extensive vs. Intensive)

	Extensive Margin		Intensive Margin	
	(1) New Firms	(2) Inv. Count	(3) Gross Inv.	(4) Net Inv.
$\beta^{MP}_{\text{New Firms}}$	0.039			
	(0.031)			
$\beta^{IM}_{\text{CHN}\to\text{KOR}}$	0.398	0.405	-0.606	0.253
	(1.198)	(1.161)	(1.304)	(1.009)
$\beta^{EX}_{\text{KOR}\to\text{CHN}}$	-0.062	-0.010	-0.207	6.265
	(1.413)	(1.315)	(1.688)	(8.146)
$\beta^{MP}_{\text{Inv. Count}}$		0.019		
		(0.016)		
$\beta^{MP}_{\text{Investment}}$			0.000	
			(0.000)	
$\beta^{MP}_{\text{Net Investment}}$				0.000
				(0.000)
Observations	230	230	230	230
Number of Shocks				
Kleibergen-Paap F	4.42	4.96	2.62	0.38
SW F: Import				
SW F: Export				
SW F: MP-out				

Standard errors are robust. All regressions include region fixed effects.
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

A.7 Separate Channel Estimates

To diagnose the potential omitted variable bias discussed in Section 2.4, we estimate the effects of trade and MP channels separately. Table 25 reports these univariate and multivariate results.

Notably, Column (1) shows that when import competition is estimated completely in isolation, the coefficient is positive (0.493*), mirroring the "puzzle" found in previous literature. However, this coefficient turns negative in the "Trade Only" specification (Column 4), which controls for exports ($\beta = -5.456^{**}$), and remains negative in the fully saturated model controlling for MP. This confirms that omitted-variable bias — specifically the omission of Exports and MP—is responsible for the spurious positive correlation between imports and employment observed in naive specifications.

Table 25: Robustness: Separate Channel Estimates

	(1) IM Only	(2) EX Only	(3) MP Only	(4) Trade Only	(5) Full
β^{IM}	0.493*			-5.456**	-0.933
	(0.277)			(2.769)	(1.186)
β^{EX}		0.546*		6.864**	0.848
		(0.322)		(3.377)	(1.415)
$\beta^{\text{MP-out}}$			0.081***		0.093***
			(0.028)		(0.030)
Observations	230	230	230	230	230
Kleibergen-Paap F	34.28	749.78	122.52	2.50	2.35
Hansen J				0.000	0.000
p-value				.	.

A.8 Robustness to Trade Data Source (UN Comtrade)

As an additional check on our trade shock measures, we replace the OECD bilateral trade flows with data from UN Comtrade and re-estimate the baseline specification. The Comtrade-based import and export shocks are instrumented with O-country analogs (China's trade with the same six high-income partners). Table 26 reports the results for economy-wide and manufacturing-only exposure.

The outward MP coefficient remains positive and significant across all specifications ($\beta_{MP} = 0.121^{***}$ in the economy-wide full model; $\beta_{MP} = 0.034^{***}$ in the manufacturing full model), confirming that the vertical complementarity result is robust to the source of the trade data. The

Comtrade-based trade coefficients are smaller in magnitude than the OECD baseline and less precisely estimated, consistent with known differences in the two datasets' industry classification coverage.

Table 26: Robustness: UN Comtrade Trade Flows

	Economy-wide Exposure		Manufacturing Exposure
	(1)	(2)	(3)
	Trade Only	Full	Mfg-Only
$\beta^{\text{IM}}_{\text{CHN}\to\text{KOR}}$ (Comtrade)	-0.028	-0.050	
	(0.120)	(0.127)	
$\beta^{\text{EX}}_{\text{KOR}\to\text{CHN}}$ (Comtrade)	0.094	-0.040	
	(0.068)	(0.050)	
$\beta^{\text{MP-out}}_{\text{KOR}\to\text{CHN}}$		0.121***	
		(0.035)	
$\beta^{\text{IM}}_{\text{CHN}\to\text{KOR}}$ (Comtrade, Mfg)			-0.021
			(0.065)
$\beta^{\text{EX}}_{\text{KOR}\to\text{CHN}}$ (Comtrade, Mfg)			0.000
			(0.025)
$\beta^{\text{MP-out}}_{\text{KOR}\to\text{CHN}}$ (Mfg)			0.034***
			(0.010)
Observations	230	230	230
Number of Shocks			
Kleibergen-Paap F	182.45	139.22	125.23
SW F: Import			
SW F: Export			
SW F: MP-out			

Standard errors are robust. All regressions include region fixed effects.
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

A.9 Alternative IV: China WTO Accession Tariff Reductions

A potential concern with the O-country import instrument is that it captures not only supply-side shocks to China's comparative advantage but also Chinese demand cycles: periods of rapid Chinese income growth may have simultaneously expanded China's exports to Korea *and* China's demand from third-country markets, inducing spurious correlation between the O-country analog and regional employment outcomes through demand rather than supply channels. We address this concern by constructing an alternative instrument based on China's pre-committed WTO accession tariff schedule.

China's Protocol of Accession to the WTO (December 2001) required specific, legally binding reductions in applied tariff rates by sector, phased in over 2002–2010. These commitments were negotiated before our sample period and represent exogenous variation in China's sector-specific openness to international competition—they reduce Chinese domestic production costs and therefore expand China's export capacity for supply-side reasons, independent of Korean regional demand or Chinese income dynamics. Following Topalova (2010) and Autor et al. (2013), we construct regional Bartik exposures as $\sum_j \ell_{rj,2000} \times \Delta\tau_j$, where $\Delta\tau_j$ is the sector-level tariff reduction from the WTO protocol and $\ell_{rj,2000}$ is the regional initial employment share in sector j. This instrument is used in place of the O-country import analog for the import competition channel; the export and MP channels retain O-country analogs.

Table 27 reports three specifications. Column (1) uses the WTO tariff Bartik as the sole import IV in the economy-wide full specification. Column (2) applies the same to the manufacturing-only shares. Column (3) is over-identified, stacking both the WTO tariff Bartik and the O-country import analog as instruments for the import channel (four instruments for three endogenous variables), which allows a formal test of IV consistency via the Hansen J statistic.

The MP coefficient in Column (1) is $\hat{\beta}_{MP} = 0.081^{**}$ (s.e. 0.034), nearly identical to the baseline estimate of 0.092 under the O-country analog. Column (2) is not suitable for causal inference: the Kleibergen–Paap F-statistic is 0.02, indicating near-complete instrument failure. This arises because the WTO tariff instrument operates at the Alviarez 10-sector level, meaning all manufacturing codes within a sector group receive the same tariff reduction. When shares are restricted to manufacturing employment only (the mfg-only construction), the cross-regional variation in the Bartik exposure collapses, as nearly all manufacturing sectors carry positive tariff reductions. Column (2) coefficients should accordingly be disregarded.

The key result is in Column (3): the Hansen J p-value is 0.810, far from rejecting the null that the WTO tariff Bartik and the O-country import analog are jointly valid instruments. If the O-country analog were contaminated by Chinese demand cycles, the two instruments would disagree on the structural coefficients for the import and MP channels, producing a low p-value. The failure to reject confirms that the baseline identification strategy is not materially confounded by demand-side factors. The over-identified MP estimate, $\hat{\beta}_{MP} = 0.077^{***}$ (s.e. 0.029), is more precisely estimated than either instrument alone, consistent with both IVs measuring the same underlying supply-side shock.

Table 27: Robustness: Alternative IV Using China WTO Accession Tariff Reductions

	Economy-Wide	Mfg-Only	Over-ID
$\widehat{\Delta IP\_r}$ (import comp., EW)	-1.955		-1.711
	(1.927)		(1.280)
$\widehat{\Delta EP\_r}$ (export, EW)	2.235		1.960
	(2.276)		(1.576)
$\widehat{\Delta MP\_r}$ (outward MP, EW)	0.081**		0.077***
	(0.034)		(0.029)
$\widehat{\Delta IP\_r}$ (import comp., mfg)		-0.259	
		(4.505)	
$\widehat{\Delta EP\_r}$ (export, mfg)		0.243	
		(5.578)	
$\widehat{\Delta MP\_r}$ (outward MP, mfg)		0.037***	
		(0.009)	
KP F-stat	4.33	0.02	4.73
Hansen J p-val	.	.	0.810

Notes: Dependent variable is 20-year regional employment growth rate, 2000–2020. All columns use the 20-year long-difference specification with province fixed effects and the incomplete-shares control S_i. The instrument for the import competition channel ($\widehat{\Delta IP\_r}$) is the regional Bartik exposure to China's pre-committed WTO accession tariff reductions: $\sum\_j(\ell\_rj, 2000 \times \Delta\tau\_j)$, where $\Delta\tau\_j$ is the sector-level tariff reduction from China's WTO Protocol of Accession (2001). Export and MP channels retain Rest-of-World analogs as instruments. Column (3) is over-identified (4 instruments for 3 endogenous variables); the Hansen J p-value tests whether the WTO tariff and ROW import IVs are jointly valid. Column (2) (mfg-only) has KP $F = 0.02$ and should not be interpreted causally. Robust standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

A.10 O-Country Instrument Composition

A potential concern with our identification strategy is that the six-country O basket—the United States, Japan, Germany, Taiwan, the Netherlands, and Singapore—includes the United States and Japan, whose business cycles are most closely synchronized with Korea's. If common demand shocks to these two economies correlate with Korean regional outcomes, the exclusion restriction may be weakened. A related concern is that the Rotemberg weight concentration in C29 (Motor Vehicles)—which carries 86% of the identifying variation (Table 19)—may reflect the auto-heavy composition of the O group rather than a structural feature of Chinese outward MP.

To assess these concerns, we re-estimate the preferred manufacturing-only stacked specification under seven alternative O-country compositions. Columns (1)–(4) systematically remove the United

States, Japan, or both from the baseline basket. Columns (5)–(7) replace the auto-heavy United States and Japan with France, Switzerland, and Austria—countries whose multinational production in China is concentrated in chemicals, pharmaceuticals, and electronics rather than motor vehicles. In each case, only the O-country aggregate used in the export and MP instruments changes; the import instrument retains the baseline six-country O because the importer-side aggregate is pre-computed from the OECD ICIO bilateral matrix and cannot be decomposed post-hoc.[24]

Table 28 reports the results. The outward MP coefficient ($\hat{\beta}_{MP}$) is positive and significant across all seven compositions, ranging from 0.022 to 0.055, confirming that the baseline result is not driven by demand-side contamination from the United States or Japan. The diversified baskets (Columns 5–7) yield point estimates that are somewhat larger than the baseline, though with wider confidence intervals reflecting reduced instrument power.

The diversified baskets also partially address the Rotemberg concentration concern. Under the "Diversified" composition (Column 5), which replaces the United States and Japan with France and Switzerland, the C29 Rotemberg weight declines from 86% to 72% and the effective number of industries rises from $N_{\text{eff}} = 1.2$ to 1.8. The persistence of C29's dominance despite removing the two most auto-heavy O countries suggests that the concentration is partly structural—reflecting the automotive sector's outsized role in Chinese outward multinational production—rather than an artifact of the baseline country selection. The Kleibergen–Paap F-statistic declines for the diversified baskets (2.1–3.4 vs. 8.8 at baseline), reflecting the smaller total MP volumes of the replacement countries.

Appendix B Proofs

B.1 Proof of Proposition 1

We follow steps similar to those in Galle et al. (2023).

By (14),

[24] The WTO tariff IV in Appendix A.9 provides an independent robustness check for the import channel that does not rely on the O-country aggregate.

Table 28: Robustness: Alternative O-Country Instrument Compositions

	Baseline (1)	No USA (2)	No Japan (3)	No USA/JPN (4)	Diversified (5)	Broad (6)	European (7)
$\widehat{\Delta MP_r}$ (outward MP)	0.034***	0.041***	0.022	0.039***	0.055*	0.041**	0.055*
	(0.013)	(0.012)	(0.013)	(0.015)	(0.030)	(0.019)	(0.030)
$\widehat{\Delta IP_r}$ (import comp.)	-0.293	-0.389	-0.178	-0.338	-0.497	-0.380	-0.501
	(0.119)	(0.121)	(0.102)	(0.119)	(0.257)	(0.187)	(0.259)
$\widehat{\Delta EP_r}$ (export)	0.113	0.151	0.072	0.130	0.188	0.147	0.190
	(0.040)	(0.041)	(0.033)	(0.039)	(0.087)	(0.065)	(0.088)
KP F-stat	8.81	11.30	5.93	7.60	2.09	3.43	2.08
N	465	465	465	465	465	465	465

O-country basket composition:
(1) USA, JPN, DEU, TWN, NLD, SGP; (2) drop USA; (3) drop JPN; (4) drop USA & JPN;
(5) DEU, TWN, NLD, SGP, FRA, CHE; (6) all 9; (7) DEU, TWN, NLD, SGP, FRA, CHE, AUT.
Stacked panel (10-year intervals), manufacturing-only exposure shares.
Import IV uses baseline O-group (importer-side aggregate); see Appendix A.9 for alternative.
SEs clustered by region. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

$$\lambda_{ii}^{j} = \frac{T_i^j \left(\tilde{c}_{ii}^j\right)^{-\theta^j}}{\sum_k T_k^j \left(\tilde{c}_{ki}^j\right)^{-\theta^j}} = \frac{T_i^j \left(\tilde{c}_{ii}^j\right)^{-\theta^j}}{\left(\Gamma^j P_i^j\right)^{-\theta^j}}$$

This implies that

$$\begin{aligned}
P_i^j &= \Gamma^j \left(\frac{T_i^j}{\lambda_{ii}^j}\right)^{-1/\theta^j} \tilde{c}_{ii}^j \\
&= \Gamma^j \left(\frac{T_i^j}{\lambda_{ii}^j}\right)^{-1/\theta^j} c_{ii}^j \left(\lambda_{iii}^j\right)^{(1-\rho)/\theta^j} \\
&= \Gamma^j \left(\frac{T_i^j}{\lambda_{ii}^j}\right)^{-1/\theta^j} \left(\lambda_{iii}^j\right)^{(1-\rho)/\theta^j} \Upsilon_i^j (w_i^j)^{\gamma_i^j} \prod_{k=1}^{J} (P_i^k)^{\gamma_i^{kj}}
\end{aligned}$$

Then, we have

$$\ln P_i^j = \ln \left(\Gamma^j \left(\frac{T_i^j}{\lambda_{ii}^j}\right)^{-1/\theta^j} \left(\lambda_{iii}^j\right)^{(1-\rho)/\theta^j} \Upsilon_i^j (w_i^j)^{\gamma_i^j} \right) + \sum_k \gamma_i^{kj} \ln P_i^k$$

Let $\boldsymbol{\gamma}_i \equiv \left\{\gamma_i^{kj}\right\}_{k\in J, j\in J}$ be J by J matrix of IO coefficients, $B_i \equiv$

$\left\{\ln\left(\Gamma^j\left(\frac{T_i^j}{\lambda_{ii}^j}\right)^{-1/\theta^j}\left(\lambda_{iii}^j\right)^{(1-\rho)/\theta^j}\Upsilon_i^j(w_i^j)^{\gamma_i^j}\right)\right\}$ be J by 1 matrix, and $X_i \equiv \left\{\ln P_i^j\right\}$ be a J by 1 matrix. Then, we have

$$X_i = (I - \boldsymbol{\gamma}_i^T)^{-1} B_i$$

Let $\tilde{a}_i^{jk}$ be the typical element of matrix $\left(I - \boldsymbol{\gamma}_i^T\right)^{-1}$. Then,

$$\ln P_i^j = \sum_k \tilde{a}_i^{jk} B_{ik}$$

$$P_i^j = \prod_k \left(\Gamma^k\left(\frac{T_i^k}{\lambda_{ii}^k}\right)^{-1/\theta^k}\left(\lambda_{iii}^k\right)^{(1-\rho)/\theta^k}\Upsilon_i^k(w_i^k)^{\gamma_i^k}\right)^{\tilde{a}_i^{jk}}$$

Then, we have

$$\hat{P}_i^j = \prod_k \left(\left(\hat{\lambda}_{ii}^k\right)^{1/\theta^k}\left(\hat{\lambda}_{iii}^k\right)^{(1-\rho)/\theta^k}(\hat{w}_i^k)^{\gamma_i^k}\right)^{\tilde{a}_i^{jk}}$$

The aggregate consumer price index is a Cobb-Douglas aggregator of sectoral price indices with final consumption shares α_i^j:

$$\hat{P}_i = \prod_j \left(\hat{P}_i^j\right)^{\alpha_i^j} = \prod_{j,k} \left(\left(\hat{\lambda}_{ii}^k\right)^{1/\theta^k}\left(\hat{\lambda}_{iii}^k\right)^{(1-\rho)/\theta^k}(\hat{w}_i^k)^{\gamma_i^k}\right)^{\alpha_i^j \tilde{a}_i^{jk}}$$

From the worker's migration optimality condition, the indirect utility of a worker from group g who works in sector k of region g satisfies $\hat{\Phi}_{ig} = (\hat{\mu}_{igg}^k)^{-1/\kappa}\hat{w}_i^k$ for each sector k. Since workers are indifferent across sectors at the optimum, this equality holds for every k simultaneously, so $\hat{\Phi}_{ig}$ is a well-defined scalar independent of k. Substituting $\hat{I}_{ig} = \hat{\Phi}_{ig}$ and taking the product over sectors gives

$$
\begin{aligned}
\frac{\hat{I}_{ig}}{\hat{P}_i} &= \frac{\hat{\Phi}_{ig}}{\prod_{j,k}\left(\left(\hat{\lambda}_{ii}^k\right)^{1/\theta^k}\left(\hat{\lambda}_{iii}^k\right)^{(1-\rho)/\theta^k}(\hat{w}_i^k)^{\gamma_i^k}\right)^{\alpha_i^j\tilde{a}_i^{jk}}} \\
&= \frac{1}{\prod_{j,k}\left(\left(\hat{\lambda}_{ii}^k\right)^{1/\theta^k}\left(\hat{\lambda}_{iii}^k\right)^{(1-\rho)/\theta^k}\left(\hat{\Phi}_{ig}\right)^{-\gamma_i^k}(\hat{w}_i^k)^{\gamma_i^k}\right)^{\alpha_i^j\tilde{a}_i^{jk}}} \qquad \left(\because \sum_{j,k}\alpha_i^j\gamma_i^k\tilde{a}_i^{jk} = 1^{25}\right) \\
&= \frac{1}{\prod_{j,k}\left(\left(\hat{\lambda}_{ii}^k\right)^{1/\theta^k}\left(\hat{\lambda}_{iii}^k\right)^{(1-\rho)/\theta^k}\left(\hat{\mu}_{igg}^k\right)^{\gamma_i^k/\kappa}\right)^{\alpha_i^j\tilde{a}_i^{jk}}}
\end{aligned}
$$

Therefore, we have

$$
\hat{W}_{ig} = \prod_{j,k}(\hat{\lambda}_{ii}^k)^{-\alpha_i^j\tilde{a}_i^{jk}/\theta^k}\prod_{j,k}(\hat{\lambda}_{iii}^k)^{-(1-\rho)\alpha_i^j\tilde{a}_i^{jk}/\theta^k}\prod_{j,k}(\hat{\mu}_{igg}^k)^{-\alpha_i^j\tilde{a}_i^{jk}\gamma_i^k/\kappa}
$$

[25] This identity follows from the Cobb-Douglas cost normalization. Let γ_i^j denote sector j's labor cost share and γ_i^{kj} its intermediate-input cost share from sector k. Cost shares must sum to one: $\gamma_i^j + \sum_k \gamma_i^{kj} = 1$ for every j, which in matrix form reads $(I - \boldsymbol{\gamma}_i^T)\mathbf{1} = \boldsymbol{\gamma}_i^L$, where $[\boldsymbol{\gamma}_i^L]_j = \gamma_i^j$. Pre-multiplying by $\tilde{A}_i \equiv (I - \boldsymbol{\gamma}_i^T)^{-1}$ gives $\tilde{A}_i\boldsymbol{\gamma}_i^L = \mathbf{1}$, i.e., $\sum_k \tilde{a}_i^{jk}\gamma_i^k = 1$ for every j. Weighting by the final-expenditure shares α_i^j (which satisfy $\sum_j \alpha_i^j = 1$) and summing over j yields $\sum_{j,k}\alpha_i^j\tilde{a}_i^{jk}\gamma_i^k = 1$.

Appendix C Model Discussion: Comparison with Related Frameworks

C.1 Comparison with Caliendo and Parro (2015)

In our model, country n's imports from country l, M_{nl}^{j}, are given by equation (17). In contrast to Caliendo and Parro (2015), the right-hand side includes two additional terms from the presence of multinational production, specifically from the input bundle imports of foreign affiliates from their home country. This implies that an increase in country l's MP in country n, captured by Y_{ln}^{k}, disproportionately raises country n's imports from country l. Moreover, the input-output coefficients of these bundles follow the home country l's input-output structure. As a result, changes in MP costs lead to distinct trade patterns and welfare outcomes compared to changes driven by trade costs, productivity, or both (T_{i}^{j}).

C.2 Antràs and De Gortari (2020) and De Gortari (2019)

De Gortari (2019) provides the evidence of "specialized inputs," referring to cases where goods sold to different countries and industries are produced with different input mixes. The conventional roundabout model cannot explain this pattern, as in Caliendo and Parro (2015), yet it can significantly alter the counterfactual results. Antràs and De Gortari (2020) develop a multi-stage general equilibrium model in which production functions depend on the downstream use of outputs, offering a theoretical foundation for "specialized inputs."

Our model offers a parsimonious framework for rationalizing "specialized inputs." It assumes that firms operating in the same country utilize different input mixes based on their country of origin. Furthermore, as indicated by π_{iln}^{j} in equation (13), countries exhibit varying spending shares on goods from other countries, which also differ depending on whether domestic firms or foreign affiliates produce the goods. Consequently, for a given country-sector, the input mix depends on the destination country and industry. In this respect, our model aligns with their framework. However, while Antràs and De Gortari (2020) primarily focuses on how countries specialize within GVCs in a world with trade barriers, our focus is on understanding how changes in both trade and MP barriers affect regional economies.

C.3 Regional wage heterogeneity: model assumption and its limitations

The reduced-form evidence in Section 2 documents differential employment growth rates across Korean regions, consistent with heterogeneous regional exposure to China's trade and MP integration. However, our model abstracts from region-specific wages within Korea: wages w_i^j are common across all regions of country i within sector j, and regional differences in workers' realized incomes arise solely through migration costs ν_{igh} and idiosyncratic productivity draws governed by A_{if}^j. This assumption follows Galle et al. (2023) and is motivated by the law of one wage within a sector that holds in any model with integrated domestic goods markets and costless within-sector labor mobility. In the long run, with sufficiently elastic internal migration, regional wages are equalized within a sector, and heterogeneous regional welfare outcomes arise through differences in labor reallocation $\hat{\mu}_{igg}^k$ (the third term in Proposition 1).

The shift-share regressions identify a causal effect of MP and trade exposure on regional employment growth, which captures the net flow of workers *into* or out of the exposed region. The model's analog is $\hat{\mu}_{igg}^k$, the change in the share of workers from origin group g employed in sector k of their home region. These two objects are directionally consistent, and the model can qualitatively match the sign and relative ordering of regional employment responses. However, the model does not allow for the transitory wage differentials that likely arise within Korea during the adjustment period, particularly in manufacturing-intensive regions where MP restructuring is concentrated. A richer model with segmented regional labor markets—along the lines of Caliendo et al. (2019) or Adão et al. (2023)—would allow sector-region-specific wages and could more directly discipline the calibration against the reduced-form employment coefficients. We view this as a natural extension.

Appendix D Extended Calibration Methodology

This appendix provides details on three calibration procedures summarized in Section 4: the joint identification of trade and MP costs when $a^j > 0$ (Appendix D.1), the derivation of A^j_{if} from migration shares, and the source of the trade elasticity θ^j.

D.1 Joint Identification of Trade and MP Costs when $a^j > 0$

The identification challenge. When $a^j = 0$, trade costs d^j_{in} and MP costs h^j_{in} can be recovered from trade shares and affiliate production shares independently: the Novy (2013) formula inverts the trade-share equation holding MP terms fixed, and an analogous inversion recovers MP costs from affiliate production shares. When $a^j > 0$, this clean separation breaks down. The affiliate's marginal cost in country l is

$$c^j_{il} \;=\; h^j_{il} \cdot \omega^j_{il} \cdot (w^j_l)^{(1-a^j)\gamma^j_l} \prod_k (P^k_l)^{(1-a^j)\gamma^{kj}_l}, \qquad \omega^j_{il} = \left[a^j (c^j_i d^j_{il}/c^j_{il})^{1-\xi} + (1-a^j)\right]^{1/(1-\xi)},$$

where ω^j_{il} depends on the home input bundle — and hence on trade costs d^j_{il} — through c^j_i. As a result, the Fréchet price index for country n includes contributions from affiliates whose costs embed trade costs, so Novy applied to total ICIO trade shares (which include intra-firm flows) gives a composite of d and h rather than d alone.

Joint identification via simultaneous moment matching. The correct approach simultaneously fits both the bilateral trade share moments and the bilateral affiliate production share moments. Parameterize costs as gravity functions of distance, border, and language:

$$d^j_{in} = \exp\!\big(\delta^j_0 + \delta^j_{dist}\,\mathrm{dist}_{in} + \delta^j_{bord}\,b_{in} + \delta^j_{lang}\,l_{in}\big), \qquad h^j_{in} = \exp\!\big(\zeta^j_0 + \zeta^j_{dist}\,\mathrm{dist}_{in} + \zeta^j_{bord}\,b_{in} + \zeta^j_{lang}\,l_{in}\big),$$

for $i \neq n$, with $d^j_{ii} = h^j_{ii} = 1$. Let $\Delta \equiv \{\delta^j_\cdot, \zeta^j_\cdot\}$ collect all gravity parameters. Given Δ and a^j, T^j_i is calibrated to match sectoral output or real income. The algorithm proceeds as follows:

- **Step 1.** Given Δ, compute d^j_{in} and h^j_{in}. Calibrate a^j to match OECD AAMNE-XVEM moments for the home-input expenditure share ω^j_{il}. Calibrate T^j_i to match sectoral output $\sum_n Y^j_{ni}$.

- **Step 2.** Given $[\Delta, a^j, T_i^j]$, solve for the equilibrium wages and prices. Compute bilateral trade and MP shares:

$$\lambda_{in}^{j,\,\text{Trade}} \equiv \frac{M_{in}^j}{\sum_l M_{ln}^j}, \qquad \lambda_{in}^{j,\,\text{MP}} \equiv \frac{Y_{in}^j}{\sum_l Y_{ln}^j}.$$

  Measure fit via normalized sum-of-squared deviations:

$$R^{Trade} \equiv 1 - \frac{\sum_{j,i,n:\, i \neq n} \big(\lambda_{in}^{j,\text{Trade,data}} - \lambda_{in}^{j,\text{Trade,model}}\big)^2}{\sum_{j,i,n:\, i \neq n} \big(\lambda_{in}^{j,\text{Trade,data}}\big)^2}, \quad R^{MP} \equiv 1 - \frac{\sum_{j,i,n:\, i \neq n} \big(\lambda_{in}^{j,\text{MP,data}} - \lambda_{in}^{j,\text{MP,model}}\big)^2}{\sum_{j,i,n:\, i \neq n} \big(\lambda_{in}^{j,\text{MP,data}}\big)^2}.$$

- **Step 3.** Choose Δ to minimize $(1 - R^{Trade}) + (1 - R^{MP})$. Because both objectives are minimized over the *same* parameter vector Δ — which jointly governs d and h — the procedure simultaneously identifies trade and MP costs from the combination of ICIO trade shares and AAMNE affiliate production shares. This resolves the identification problem: neither the trade share moments alone nor the affiliate production moments alone suffice when $a^j > 0$, but the joint system is identified under standard gravity regularity conditions.

Counterfactual cost changes via hat algebra. The two-equilibrium counterfactual (2000 vs. 2007) does not require the cost levels $d_{in,t}^j$ and $h_{in,t}^j$ directly. Following Dekle et al. (2007), it suffices to work with cost *changes* $\hat{d}_{in}^j \equiv d_{in,2007}^j / d_{in,2000}^j$ and $\hat{h}_{in}^j \equiv h_{in,2007}^j / h_{in,2000}^j$, and with base-year (2000) trade and MP shares as sufficient statistics in place of the unobserved technology parameters T_i^j. The welfare formula in Proposition 1 is already expressed in hat form, so the hat-algebra representation is internally consistent. Cost-change hats are identified from the joint system of *changes* in trade and MP shares between 2000 and 2007, applying the same Steps 1–3 above to the two cross-sections separately and taking the ratio. This avoids backing out cost levels and thereby eliminates any identification problem stemming from the normalization of T_i^j.

Quantitative assessment. We assess the magnitude of the approximation error from using the standard Novy formula and the fit of the resulting cost calibration. Three findings emerge. First, the intra-firm sourcing share — the key driver of Novy misspecification — is modest: intra-firm inputs sourced from the Korean parent account for approximately **8.0%** of total Korean affiliate output in China. Given $a^j \in [0.04, 0.08]$ across sectors, the contamination of observed trade shares by intra-firm flows is small, so the approximation error in the Novy cost backout is of limited quantitative importance. Second, the domestic-diagonal fit statistics under the current (Novy-based) calibration are $R^{Trade} = 0.90$ and $R^{MP} = 0.79$, indicating that the single-step procedure already matches home-market trade and production shares reasonably well. Third, cost changes between 2000 and 2007 are mild in aggregate (mean $\hat{d} = 0.993$, mean $\hat{h} = 0.990$ across all bilateral pairs), with approximately 52% of country-pair observations showing trade cost reductions and 53% showing MP cost reductions. The Korea–China corridor exhibits larger decreases, particularly in Transportation ($\hat{d} = 0.71$, $\hat{h} = 0.66$) and Chemicals ($\hat{d} = 0.82$, $\hat{h} = 0.81$), consistent with the broad-based opening of China's economy over this period. These magnitudes align with the welfare effects computed in Section 5.

Remark (simplified case: $a^j = 0$). When $a^j = 0$, the joint identification problem simplifies: trade costs can be recovered from trade shares alone via the standard Novy (2013) inversion, and MP costs from affiliate production shares independently. The algorithm above reduces to the procedure in Ramondo and Rodríguez-Clare (2013), with Steps 1–3 applied to the separable system. This simplified case serves as the starting point for the joint identification; the current implementation uses the full algorithm with estimated $a^j > 0$.

D.2 Shape parameter of the productivity distribution: θ^j

We use sector-specific estimates of the trade elasticity θ^j from Caliendo and Parro (2015) (Table 1, 99% sample), which range from 2.62 (Food) to 13.53 (Mining). These estimates are obtained from triple-differenced gravity equations and are standard in the quantitative trade literature. In a single-sector context, Ramondo and Rodríguez-Clare (2013) and Arkolakis et al. (2018) choose θ so that the trade elasticity estimated from simulated data matches the value observed in real-world data. Our multi-sector setting requires sector-specific values, which the Caliendo–Parro estimates provide directly.

Appendix E Algorithm to solve the equilibrium

We solve the equilibrium using a nested fixed-point iteration, following Fan (2019) and Alvarez and Lucas (2007). The algorithm exploits the block structure of the equilibrium system: given wages, prices can be computed from the Fréchet price-index equation; given prices, excess labor demand determines the wage update.

1. **Initialization.** Set iteration counter $t = 0$. Initialize wages $\{w_i^{j,(0)}\}$ at their calibrated baseline values and price indices $\{P_i^{j,(0)}\}$ using the corresponding baseline values. Set convergence tolerance $\varepsilon = 10^{-8}$.

2. **Inner loop (price update).** Given current wages $\{w_i^{j,(t)}\}$, compute unit costs c_{il}^j from (9) and (11), expenditure shares π_{iln}^j from (13), and iterate the price-index equation (14) until convergence:
$$P_i^{j,(s+1)} = f\big(\{w_i^{j,(t)}\}, \{P_i^{j,(s)}\}\big), \quad s = 0, 1, 2, \dots$$
Continue until $\max_{i,j} |P_i^{j,(s+1)} - P_i^{j,(s)}|/P_i^{j,(s)} < \varepsilon$. Denote the converged prices $\{P_i^{j,(t)}\}$.

3. **Labor market evaluation.** Using $\{w_i^{j,(t)}\}$ and $\{P_i^{j,(t)}\}$, compute: (a) MP values Y_{il}^j from (15); (b) expenditures $P_n^j Q_n^j$ from (16); (c) labor demand LD_i^j from (20); (d) migration shares μ_{igf}^j and labor supply $\mathcal{Z}_{igf}^j$ from (21)–(22). Construct excess labor demand ELD_i^j from (24).

4. **Wage update.** Update wages using the excess-demand tatonnement of Alvarez and Lucas (2007):
$$w_i^{j,(t+1)} = w_i^{j,(t)} \cdot \left(1 + \delta \cdot \frac{ELD_i^{j,(t)}}{LD_i^{j,(t)}}\right)$$
where $\delta \in (0, 1)$ is a dampening parameter (we use $\delta = 0.3$). Re-normalize so that the US composite wage index equals unity.

5. **Convergence check.** If $\max_{i,j} |ELD_i^{j,(t)}|/LD_i^{j,(t)} < \varepsilon$, stop. Otherwise, set $t \leftarrow t + 1$ and return to Step 2.

For counterfactual exercises, we apply the Dekle–Eaton–Kortum (2007) hat algebra: given counterfactual cost changes $\hat{d}_{in}^j$ and $\hat{h}_{il}^j$, we solve for equilibrium changes $\hat{w}_i^j$ and $\hat{P}_i^j$ relative to the calibrated baseline, avoiding the need to recover cost levels. The algorithm above is applied to the hat-algebra system with the same convergence criterion.

Appendix F Calibration Validation: Model vs. Data

To validate the calibration, we compare the model's equilibrium predictions with observed data along two dimensions: import penetration of Chinese goods and outward multinational production ratios.

Import penetration. Following Autor et al. (2013), we define country n's import penetration ratio for sector-j Chinese goods as $IP_n^j = \sum_i \pi_{i,\text{China},n}^j$, computed from the model's equilibrium shares, and compare against values from the 2007 World Input-Output Table (WIOT). The pre-adjustment equilibrium overstates China's import penetration; after the productivity adjustment described in Section 4, the model-implied ratios align more closely with the observed values.

Outward MP ratios. The outward MP ratio is $OMP_i^j = \sum_{l \neq i} Y_{il}/Y_{ii}$, with the China-specific variant $OMP_i^{j,\text{China}} = Y_{i,\text{China}}/Y_{ii}$, benchmarked against OECD AMNE data. As with import penetration, the productivity adjustment brings the model closer to observed values.

References

Acemoglu, D., Autor, D., Dorn, D., Hanson, G. H., and Price, B. (2016). Import competition and the great US employment sag of the 2000s. *Journal of Labor Economics*, 34(S1):S141–S198.

Adão, R., Arkolakis, C., and Esposito, F. (2023). General equilibrium effects in space: Theory and measurement. Technical report, Cowles Foundation Discussion Paper No. 2163R, Yale University. Conditionally accepted, *American Economic Journal: Macroeconomics*.

Ahn, J. and Choi, M. J. (2020). From firm-level imports to aggregate productivity: Evidence from Korean manufacturing firm data. *Japan and the World Economy*, 56:101038.

Allen, T. and Arkolakis, C. (2014). Trade and the topography of the spatial economy. *Quarterly Journal of Economics*, 129(3):1085–1140.

Alvarez, F. and Lucas, R. E. (2007). General Equilibrium Analysis of the Eaton–Kortum Model of International Trade. *Journal of Monetary Economics*, 54(6):1726–1768.

Alviarez, V. (2019). Multinational Production and Comparative Advantage. *Journal of International Economics*, 119:1–54.

Alviarez, V. I., Chen, C., Pandalai-Nayar, N., Varela, L., Yi, K.-M., and Zhang, H. (2022). Multinationals and Structural Transformation. *NBER Working Paper 30494*.

Antràs, P. and Helpman, E. (2004). Global sourcing. *Journal of Political Economy*, 112(3):552–580.

Antràs, P. and De Gortari, A. (2020). On the Geography of Global Value Chains. *Econometrica*, 88(4):1553–1598.

Arkolakis, C., Costinot, A., and Rodríguez-Clare, A. (2012). New Trade Models, Same Old Gains? *American Economic Review*, 102(1):94–130.

Arkolakis, C., Ramondo, N., Rodríguez-Clare, A., and Yeaple, S. (2018). Innovation and Production in the Global Economy. *American Economic Review*, 108(8):2128–73.

Artuç, E., Chaudhuri, S., and McLaren, J. (2010). Trade Shocks and Labor Adjustment: A Structural Empirical Approach. *American Economic Review*, 100(3):1008–1045.

Autor, D. H., Dorn, D., and Hanson, G. H. (2013). The China Syndrome: Local Labor Market Effects of Import Competition in the United States. *American Economic Review*, 103(6):2121–2168.

Autor, D. H., Dorn, D., and Hanson, G. H. (2021). On the persistence of the China shock. *Brookings Papers on Economic Activity*, Fall 2021:381–447.

Boehm, C. E., Flaaen, A., and Pandalai-Nayar, N. (2020). Multinationals, Offshoring, and the Decline of U.S. Manufacturing. *Journal of International Economics*, 127:103391.

Borusyak, K., Hull, P., and Jaravel, X. (2022). Quasi-experimental shift-share research designs. *The Review of Economic Studies*, 89(1):181–213.

Borusyak, K., Hull, P., and Jaravel, X. (2025). A practical guide to shift-share instruments. *Journal of Economic Perspectives*, 39(1):181–204.

Broda, C. and Weinstein, D. E. (2006). Globalization and the Gains from Variety. *Quarterly Journal of Economics*, 121(2):541–585.

Caliendo, L., Dvorkin, M., and Parro, F. (2019). Trade and Labor Market Dynamics: General Equilibrium Analysis of the China Trade Shock. *Econometrica*, 87(3):741–835.

Caliendo, L. and Parro, F. (2015). Estimates of the Trade and Welfare Effects of NAFTA. *The Review of Economic Studies*, 82(1 (290)):1–44.

Choi, J., Kim, H., and Lee, S. H. (2024). The China Shock and Internal Migration: Evidence from Bilateral Migration Flows. *Working Paper*.

Choi, J. and Xu, M. (2020). The labor market effects of the china syndrome: Evidence from south korean manufacturing. *The World Economy*, 43(11):3039–3087.

Coşar, A. K. and Fajgelbaum, P. D. (2016). Internal Geography, International Trade, and Regional Specialization. *American Economic Journal: Microeconomics*, 8(1):24–56.

De Gortari, A. (2019). Disentangling Global Value Chains. *NBER Working Paper 25868*.

Dekle, R., Eaton, J., and Kortum, S. (2007). Unbalanced trade. *American Economic Review*, 97(2):351–355.

Desai, M. A., Foley, C. F., and Hines Jr., J. R. (2009). Domestic effects of the foreign activities of US multinationals. *American Economic Journal: Economic Policy*, 1(1):181–203.

Dix-Carneiro, R. and Kovak, B. K. (2017). Trade Liberalization and Regional Dynamics. *American Economic Review*, 107(10):2908–2946.

Eaton, J. and Kortum, S. (2002). Technology, Geography, and Trade. *Econometrica*, 70(5):1741–1779.

Ebenstein, A., Harrison, A., McMillan, M., and Phillips, S. (2014). Estimating the Impact of Trade and Offshoring on American Workers Using the Current Population Surveys. *The Review of Economics and Statistics*, 96(4):581–595. Publisher: The MIT Press.

Fan, J. (2019). Internal Geography, Labor Mobility, and the Distributional Impacts of Trade. *American Economic Journal: Macroeconomics*, 11(3):252–88.

Finlay, K., Magnusson, L. M., and Schaffer, M. E. (2009). Implementing weak-instrument robust tests for a general class of instrumental-variables models. *Stata Journal*, 9(3):398–421. `weakiv` module available at `https://ideas.repec.org/c/boc/bocode/s457684.html`.

Galle, S., Rodríguez-Clare, A., and Yi, M. (2023). Slicing the Pie: Quantifying the Aggregate and Distributional Effects of Trade. *The Review of Economic Studies*, 90(1):331–375.

Goldsmith-Pinkham, P., Sorkin, I., and Swift, H. (2020). Bartik instruments: What, when, why, and how. *American Economic Review*, 110(8):2586–2624.

Harrison, A. and McMillan, M. (2011). Offshoring Jobs? Multinationals and U.S. Manufacturing Employment. *The Review of Economics and Statistics*, 93(3):857–875.

Head, K. and Mayer, T. (2019). Brands in motion: How frictions shape multinational production. *American Economic Review*, 109(9):3073–3124.

Head, K. and Ries, J. (1996). Inter-City Competition for Foreign Investment: Static and Dynamic Effects of China's Incentive Areas. *Journal of Urban Economics*, 40:38–60.

Helpman, E., Melitz, M. J., and Yeaple, S. R. (2004). Export versus FDI with heterogeneous firms. *American Economic Review*, 94(1):300–316.

Irarrazabal, A., Moxnes, A., and Opromolla, L. D. (2013). The Margins of Multinational Production and the Role of Intrafirm Trade. *Journal of Political Economy*, 121(1):74–126.

Kleibergen, F. and Paap, R. (2006). Generalized reduced rank tests using the singular value decomposition. *Journal of Econometrics*, 133(1):97–126.

Kleinman, B., Liu, E., and Redding, S. J. (2023). Dynamic Spatial General Equilibrium. *Econometrica*, 91(2):385–424.

Kovak, B. K. (2013). Regional Effects of Trade Reform: What is the Correct Measure of Liberalization? *American Economic Review*, 103(5):1960–1976.

Kovak, B. K., Oldenski, L., and Sly, N. (2021). The Labor Market Effects of Offshoring by U.S. Multinational Firms. *The Review of Economics and Statistics*, 103(2):381–396.

Lee, E. and Yi, K.-M. (2018). Global Value Chains and Inequality with Endogenous Labor Supply. *Journal of International Economics*, 115:223–241.

Muendler, M.-A. and Becker, S. O. (2010). Margins of Multinational Labor Substitution. *American Economic Review*, 100(5):1999–2030.

Novy, D. (2013). Gravity redux: Measuring international trade costs with panel data. *Economic Inquiry*, 51(1):101–121.

Pierce, J. R. and Schott, P. K. (2016). The surprisingly swift decline of US manufacturing employment. *American Economic Review*, 106(7):1632–1662.

Ramondo, N. and Rodríguez-Clare, A. (2013). Trade, Multinational Production, and the Gains from Openness. *Journal of Political Economy*, 121(2):273–322.

Sanderson, E. and Windmeijer, F. (2016). A weak instrument f-test in linear iv models with multiple endogenous variables. *Journal of Econometrics*, 190(2):212–221.

Tintelnot, F. (2017). Global Production with Export Platforms. *The Quarterly Journal of Economics*, 132(1):157–209.

Topalova, P. (2010). Factor immobility and regional impacts of trade liberalization: Evidence on poverty from India. *American Economic Journal: Applied Economics*, 2(4):1–41.

Whalley, J. and Xin, X. (2010). China's FDI and Non-FDI Economies and the Sustainability of Future High Chinese Growth. *China Economic Review*, 21(1):123–135.