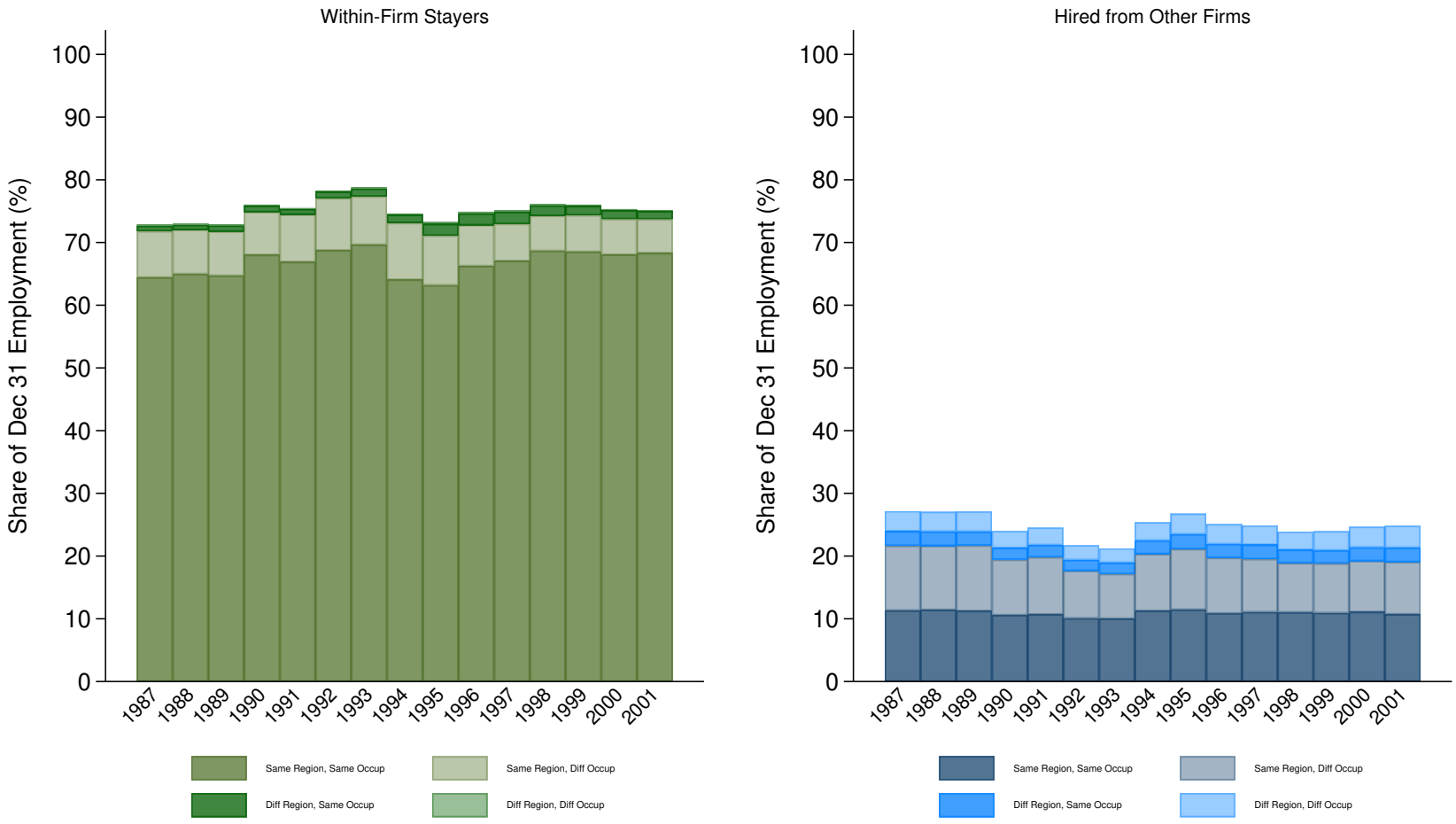
Online Appendix for

"Trade, Labor Market Concentration, and Wages,"

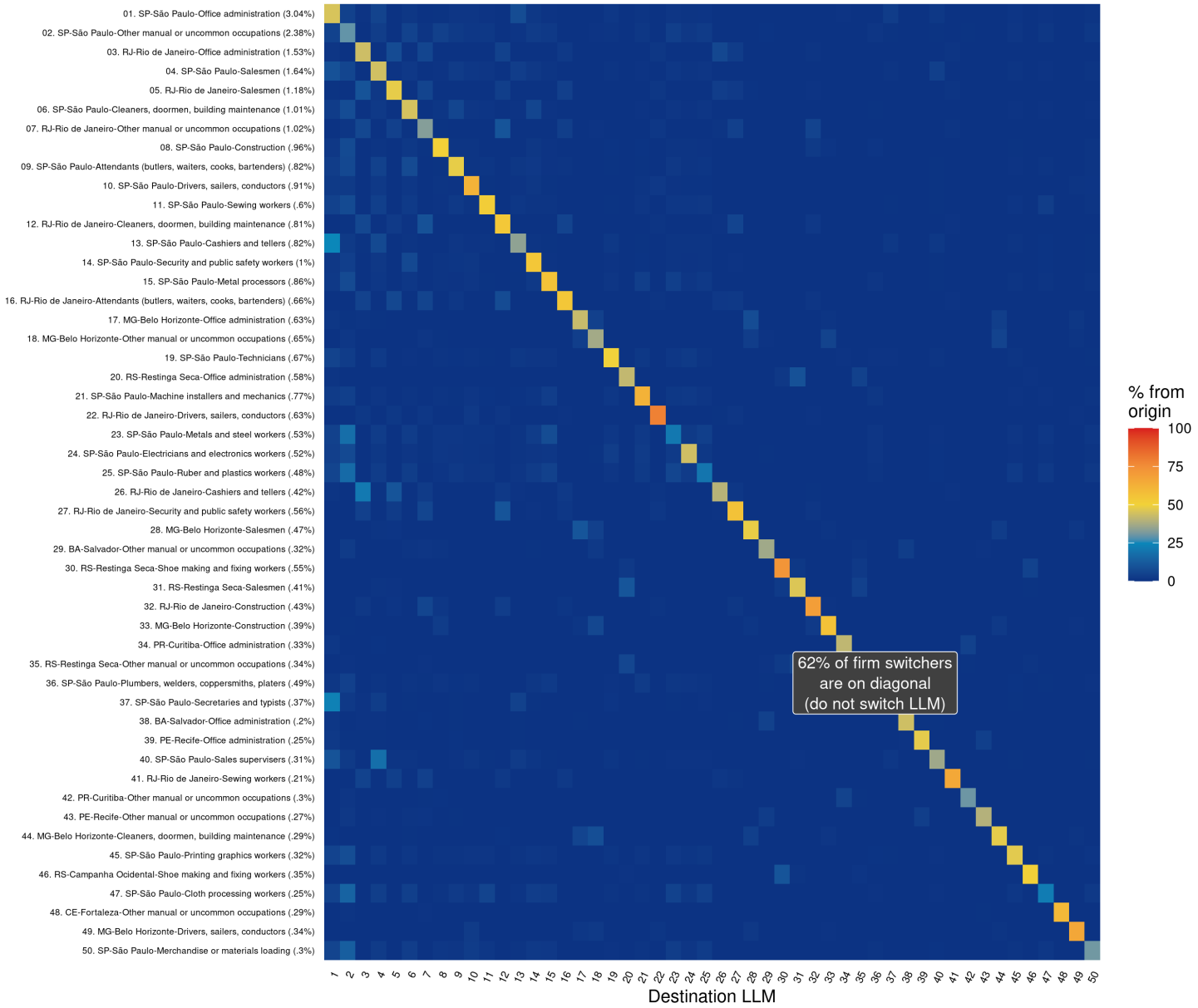
by Mayara Felix

Figure A.1: Region and occupation origins of firm stayers and firm switchers as of Dec 31 of each year



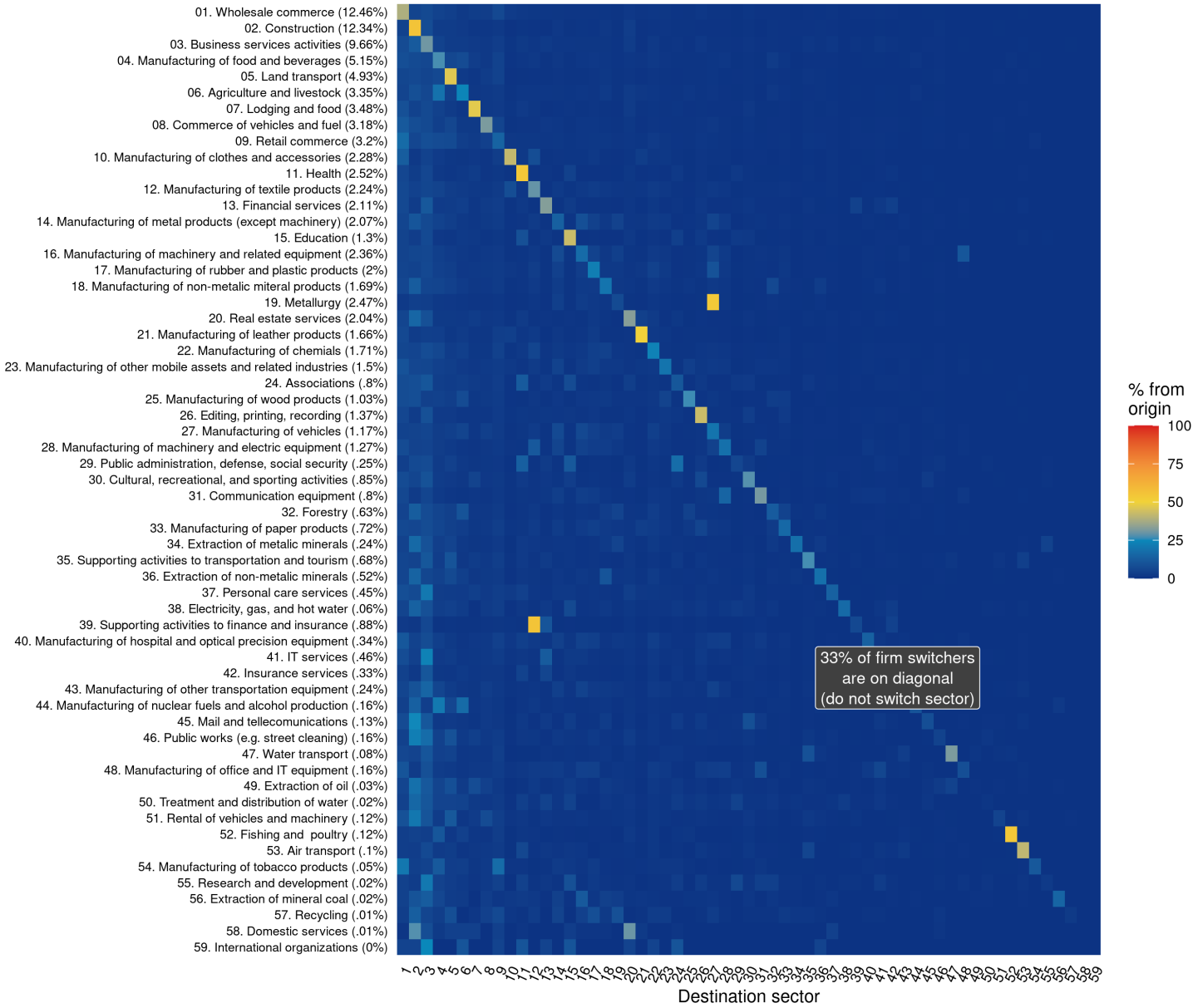
Notes: Composition of formal private sector employment in RAIS, 1986–2000. Each bar reports the December 31 employment shares of firm stayers (workers employed at the same firm on December 31 of the previous year) and of firm switchers, by region of origin in one panel and by occupation of origin in the other. Flows exclude retirements, deaths, and transitions to and from public administration.

Figure A.2: 1990-1991 local labor market transitions conditional on switching firms (Top 50)



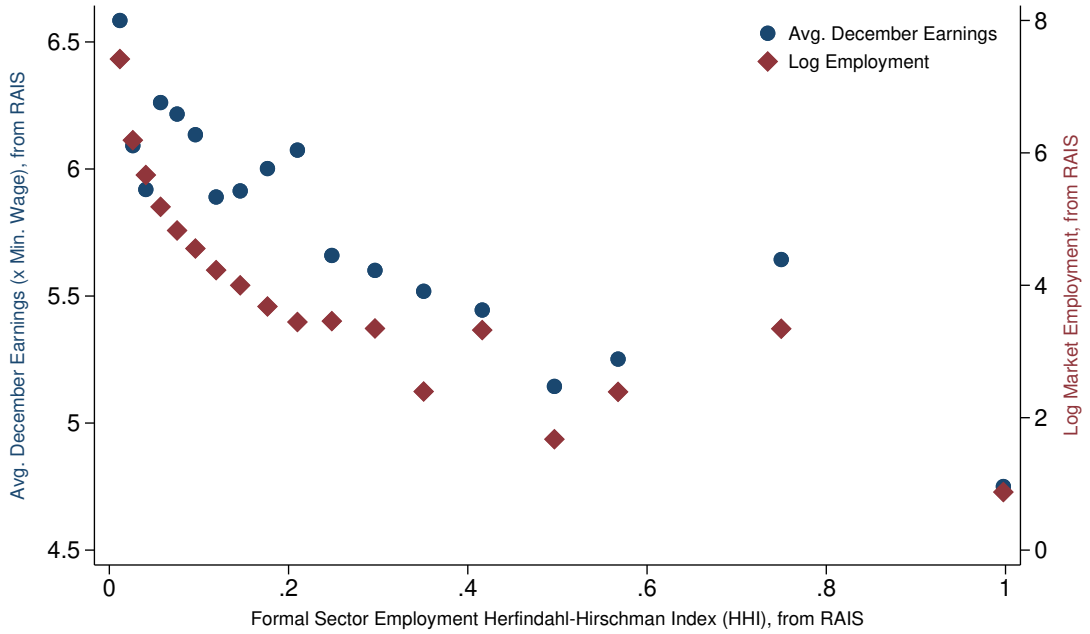
Note: This figure plots worker local labor market to local labor market transitions, among workers who switched employers between 1990 and 1991, for the top 50 local labor markets by number of workers at origin. A local labor market is a microregion × occupational group pair. Each row lists the origin microregion (with percent of total workers indicated in parentheses), while each column lists the destination microregion.

Figure A.3: 1990-1991 sector transitions conditional on switching firms (Top 50)



Note: This figure plots worker sector group to sector group transitions, among workers who switched employers between 1990 and 1991, for the top 50 sector (2-digit CNAE95) groups by number of workers at origin. Each row lists the origin sector group (with percent of total workers indicated in parentheses), while each column lists the destination sector group.

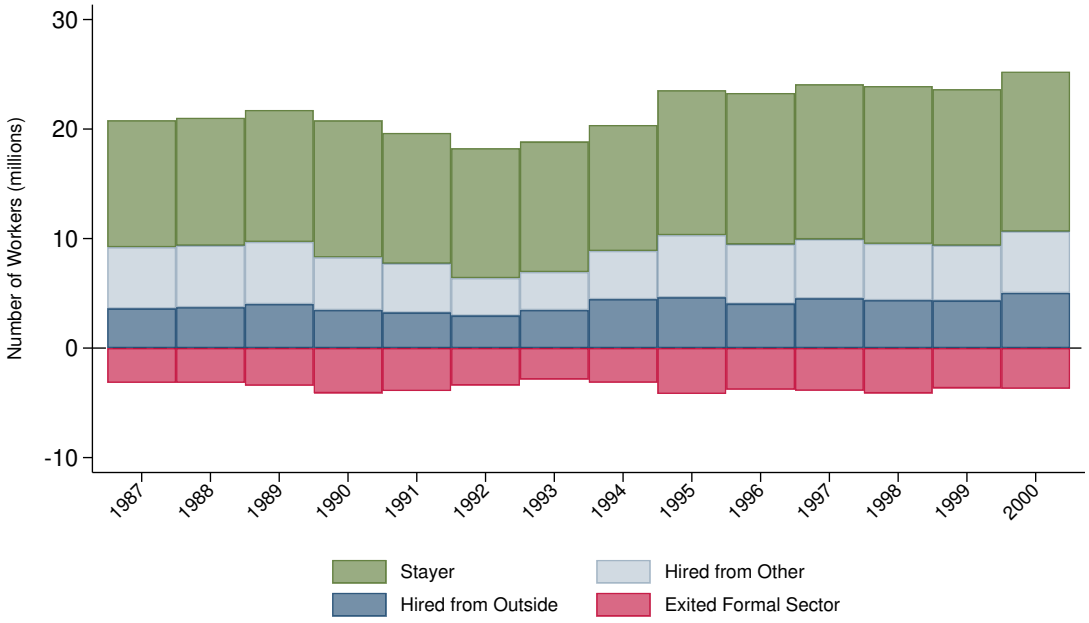
Figure A.4: 1991 wages, employment, and formal sector concentration



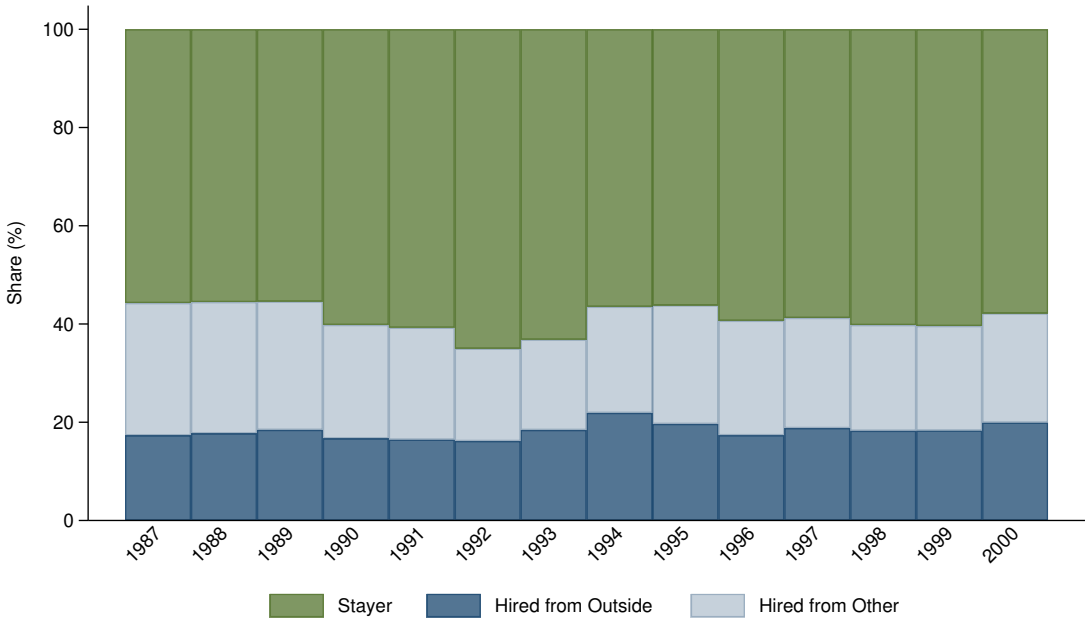
Notes: This figure plots wages and employment as a function of formal sector HHI in microregion x occupation pairs using RAIS data for 1991.

Figure A.5: Formal sector worker flows

(a) Total firm-to-firm and firm-to-outside flows



(b) Flow shares

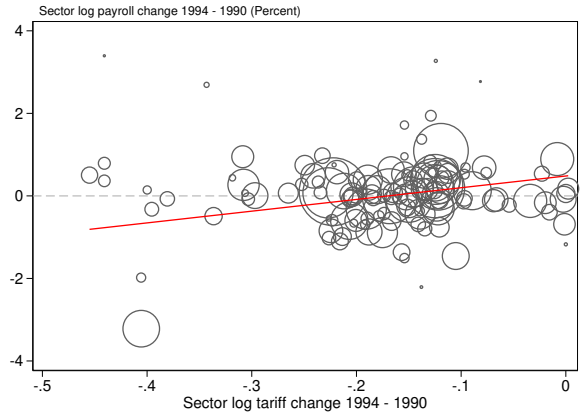


Notes: Worker flows as reported by formal sector firms in RAIS: hires directly from other formal firms ("Hired from Other"), separations followed within one year by a match to another formal firm ("Sep to Other"), hires from outside the formal sector ("Hired from Outside"), separations to outside the formal sector ("Sep to Outside"), and workers who do not switch employers within the year ("Stayer"). "Hired from Other" and "Sep to Other" are measured separately, the former from the records of the receiving firm and the latter from those of the origin firm. Flows include within-year transitions and exclude flows to and from public administration, retirements, and deaths.

Figure A.6: Changes in sector-level outcomes Before vs. After liberalization

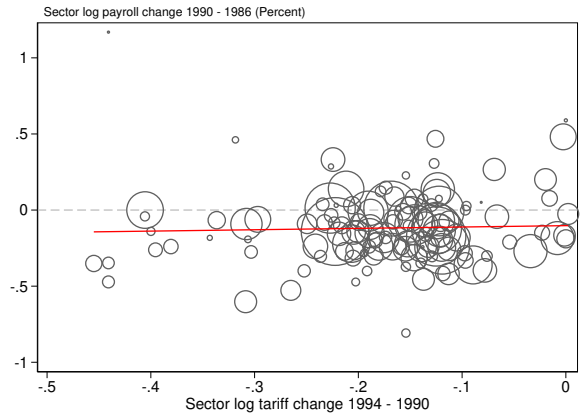
(a) After liberalization (1990–1994)





(b) Before liberalization (1986–1990)

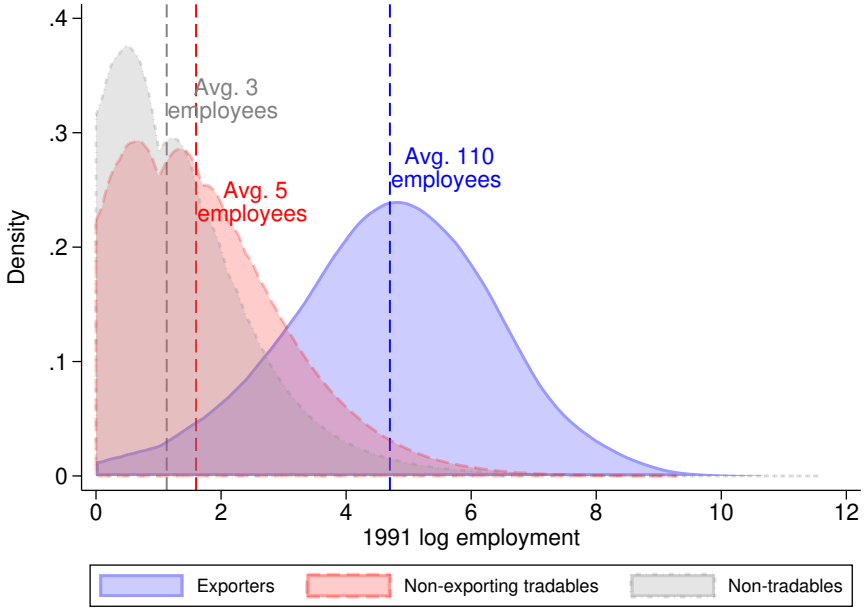




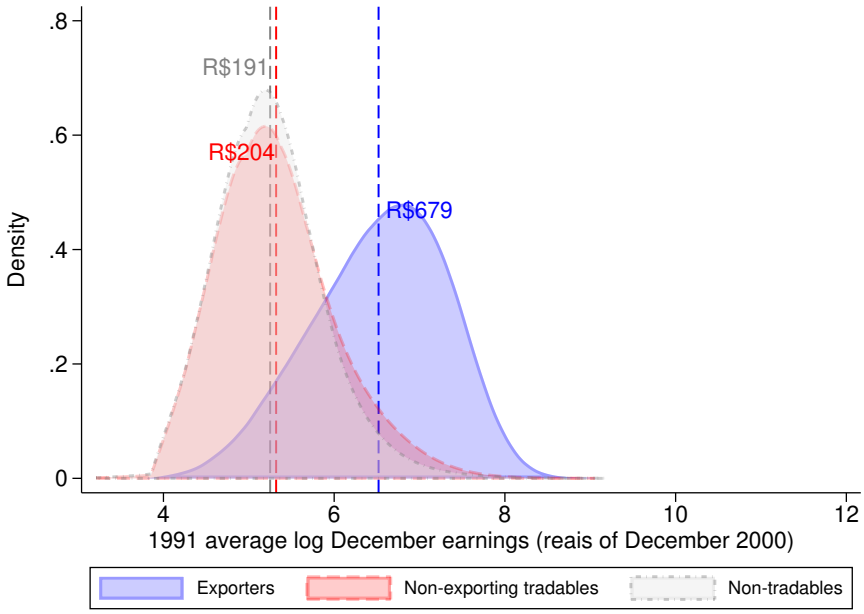
Notes: Changes in sector log employment and log payroll in the four years before and after liberalization against the 1990–1994 import tariff reduction on the sector's output. Sector totals are for the entire country, in 148 tradable sectors of RAIS's 4-digit "ibgesubatividade" variable, consistently reported throughout the period.

Figure A.7: Pre-liberalization distributions of firm size and wages by firm type

(a) Firm log employment



(b) Firm average log December earnings

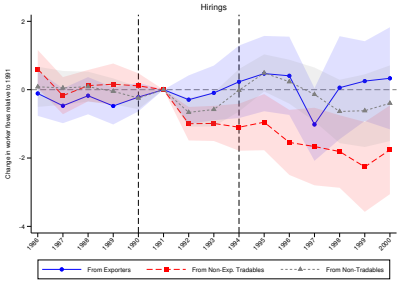


Notes: 1991 distributions of firm log employment and of firm average log December earnings (reais of December 2000) for exporters, non-exporting tradable firms, and non-tradable firms. Dashed lines mark each group's mean, annotated in levels. Exporters are firms that exported in 1991; tradable firms are in sectors with an import tariff.

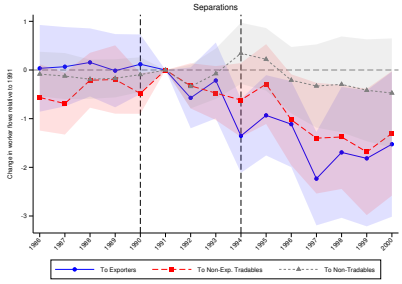
Figure A.8: Effect of import competition on worker flows into and out of non-exporting tradable firms and exporters

Within-market flows

Panel A: Hirings into non-exporting tradable firms

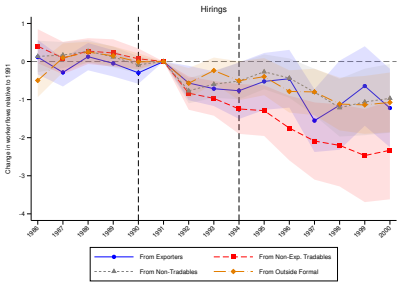


Panel B: Separations from non-exporting tradable firms

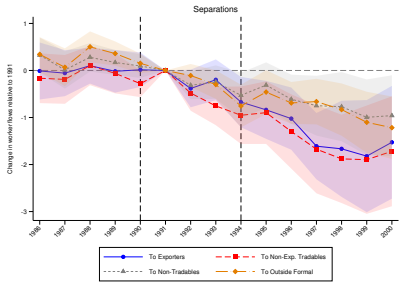


All flows

Panel A: Hirings into non-exporting tradable firms

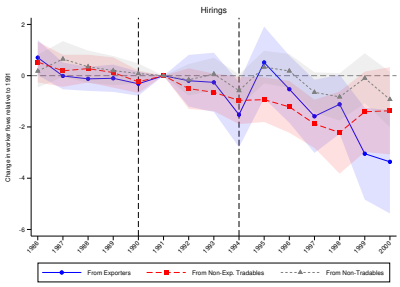


Panel B: Separations from non-exporting tradable firms

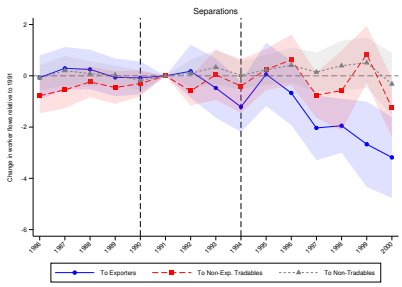


Within-market flows into and out of exporters

Panel C: Hirings into exporters



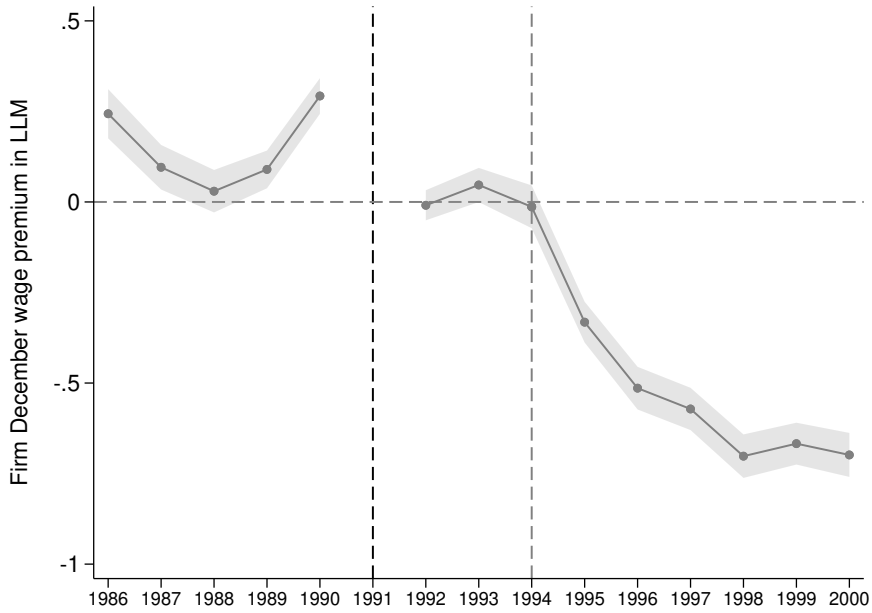
Panel D: Separations from exporters



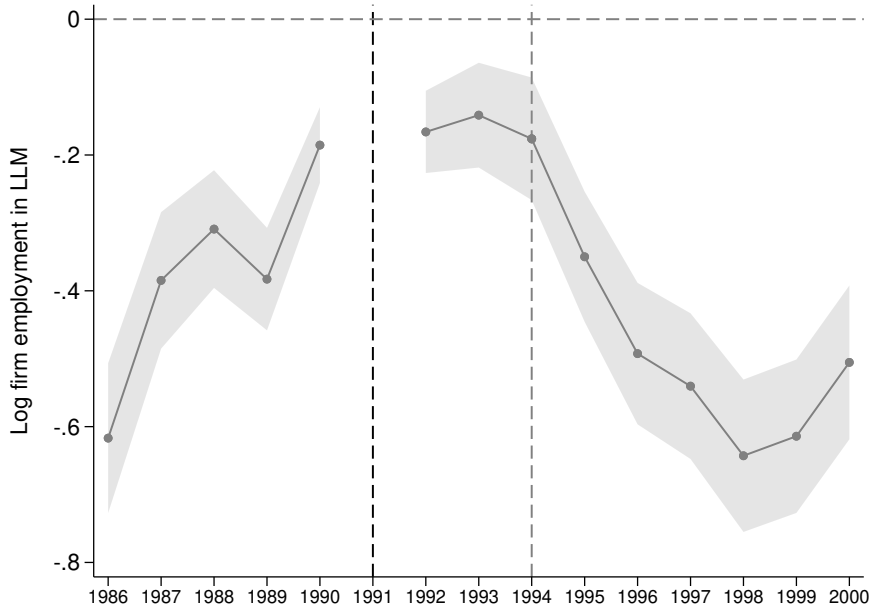
Note: This figure plots the ζ_k coefficients of equation 12 for worker flows, with local labor market (microregion × occupation) and year fixed effects and markets weighted by their 1991 formal employment, as in Table 1, column (1); standard errors are clustered by local labor market. Panel A shows hirings into non-exporting tradable firms from exporters, from other non-exporting tradable firms, and from non-tradable firms; Panel B shows separations from non-exporting tradable firms to each of these firm types. The top row restricts to flows within the same local labor market; the middle row includes all flows, including flows into and out of the formal sector. Panels C and D show within-market hirings into and separations from exporters, by the type of firm on the other side of the flow. Outcomes are inverse-hyperbolic-sine transformed counts, in long differences relative to 1991. Shaded areas are 95% confidence intervals.

Figure A.9: Year by year effects of firm-level import tariff reductions on changes in firms' local labor market wage premia and employment

(a) Wage premia

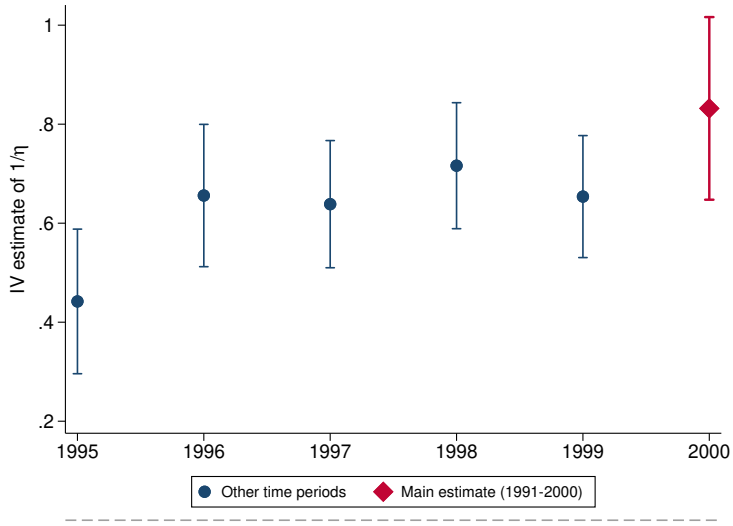


(b) Employment



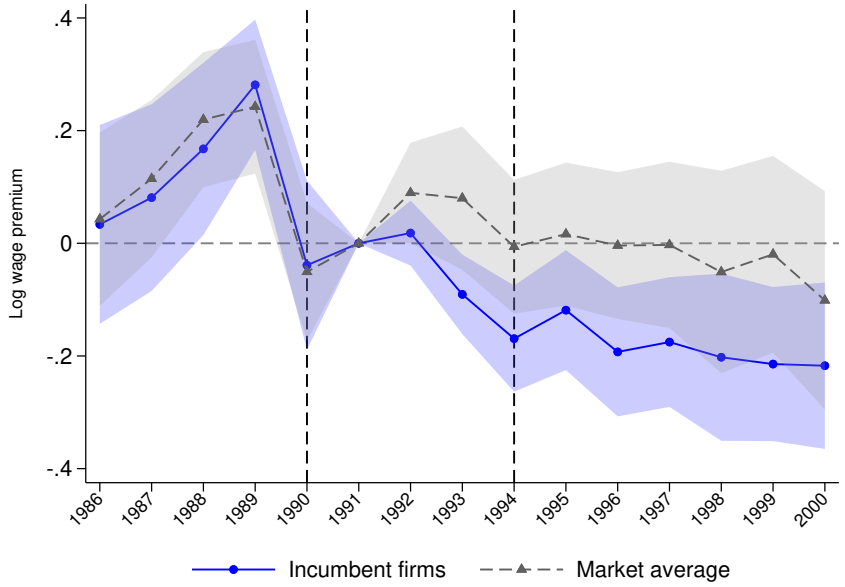
Notes: Coefficients from year-by-year regressions of the change in the firm's wage premium (Panel a) and in its log employment (Panel b) relative to 1991 on $-\ln\left(\frac{1+\tau_{1994}}{1+\tau_{1990}}\right)$, the firm's import tariff reduction. For years before 1991 the outcome is the 1991 value minus the year's value. Regressions include local labor market fixed effects and use the weights of Table 2; standard errors are clustered by firm. Dotted lines mark the beginning and end of the tariff reductions.

Figure A.10: Within-market cross-firm inverse elasticity of substitution $1/\eta$ by end year



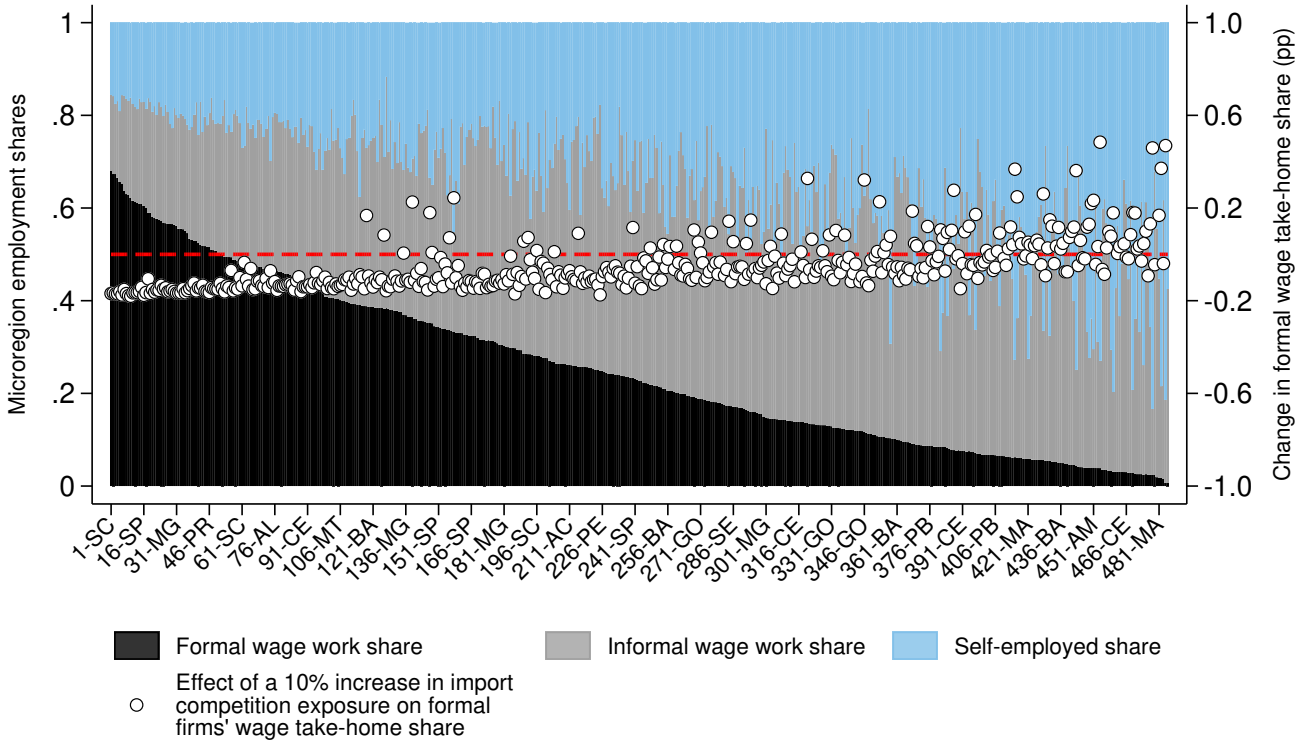
Notes: 2SLS estimates of $1/\eta$ and 95% confidence intervals from equations 15 and 16, with long differences from 1991 to each end year and the specification of Table 2, column (1). Standard errors clustered by firm.

Figure A.11: Effect of import competition on wage premia: incumbents and market average



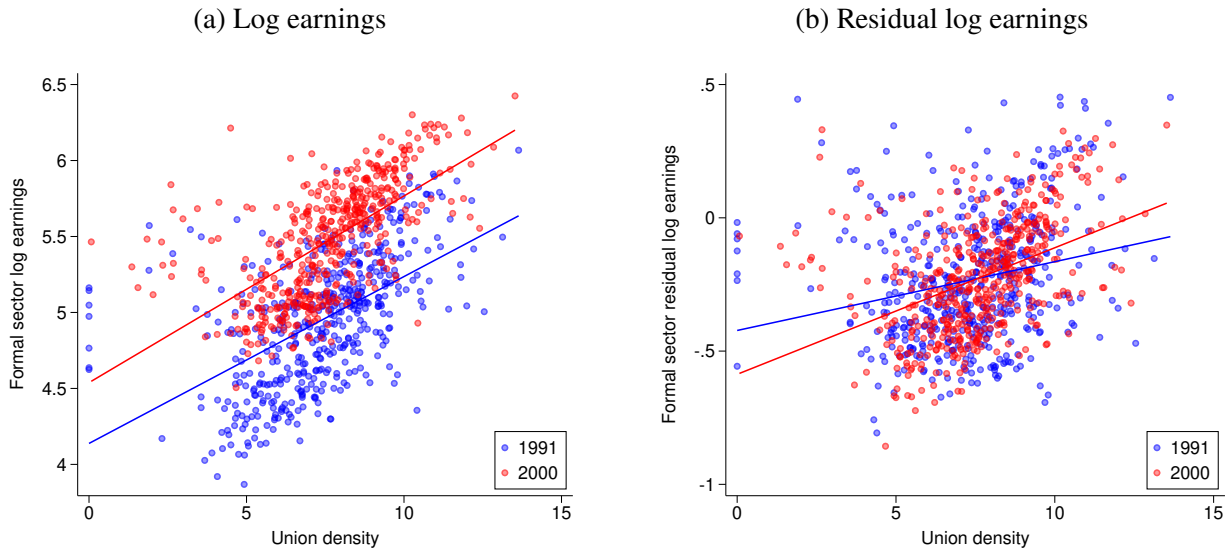
Notes: Coefficients on ΔICE_m and 95% confidence intervals from equation 12, weighted by the market's 1991 formal employment, with standard errors clustered by local labor market, in the markets of Table 5. "Incumbent firms" is the change in the wage premium of the firms present in the market in both 1991 and the year indicated, weighted by their 1991 employment. "Market average" is the change in the employment-weighted average wage premium of all of the market's firms. As in Figure 5, differences run forward in time: for the years before 1991, the outcome is its 1991 value minus its value in that year. Dashed lines indicate the beginning and end of the tariff reductions.

Figure A.12: Extended model: effect of a 10% increase in the trade shock on the formal sector wage take-home share, by microregion



Notes: As Figure 9, with every microregion receiving the same 10% increase in the trade shock ΔICE^* instead of its own 1991–2000 shock. Bars (left axis): the microregion's 1991 census shares of employment in formal wage work (black), informal wage work (gray), and self-employment (blue), for the 486 microregions of the census, sorted from the highest to the lowest formal wage work share. White circles (right axis), for the 478 microregions of the extended model: the change in the wage take-home share, in percentage points, implied by a 10% increase in ΔICE^* per equations 24 and 25, evaluated at the microregion's 1991 expected payroll Herfindahl and wage work share with the elasticities and reduced-form effects of Table 7, column (3); differences across microregions come only from their 1991 market structure. The red dashed line marks zero.

Figure A.13: Formal sector earnings and union density



Notes: Each dot is a microregion in 1991 (blue) or 2000 (red), with a linear fit for each year. Panel (a) plots the average log earnings of the microregion's formal sector workers against union density, and Panel (b) the same earnings net of observable worker characteristics. The horizontal axis is log(1 + union density). Union density data comes from https://academic.oup.com/jeea/article/23/1/236/7628306; it was collected for 1992 and 2001 and is matched to the 1991 and 2000 censuses.

Table A.1: 1991 descriptive statistics of formal local labor markets (microregion x occupation group pairs)

	Mean	Std. Dev.	p10	p25	Median	p75	p90	N
Number of formal workers	689	5,234	6	16	59	257	992	19,757
Number of formal tradable sector workers	251	1,917	0	2	15	83	346	19,757
Number of exporter workers	252	2,064	0	1	10	68	328	19,757
Number of formal non-tradable sector workers	438	3,789	2	6	27	129	533	19,757
Number of formal firms	114	790	3	6	16	54	180	19,757
Number of exporter firms	17	116	0	1	2	8	26	19,757
Payroll HHI (premium-based)	0.287	0.252	0.042	0.094	0.210	0.406	0.656	19,757
Payroll HHI (based on wage level)	0.292	0.252	0.045	0.098	0.216	0.412	0.663	19,757
Employment HHI	0.239	0.238	0.027	0.065	0.158	0.333	0.556	19,757
Average December wage (R$ of Dec. 2000)	483	495	137	193	317	571	1,019	19,757
Average December wage premium (R$ of Dec. 2000)	205	126	91	121	170	250	363	19,509
Import competition exposure (ΔICE)	0.126	0.094	0.000	0.048	0.132	0.179	0.234	19,757

Note: This table presents descriptive statistics across 19,757 Brazilian local labor markets defined as microregion × occupation group pairs, in the baseline year (1991). Means are unweighted. HHI denotes the Herfindahl-Hirschman Index of employment or payroll concentration; the premium-based Payroll HHI values each firm's payroll at its wage premium (Table 1). ΔICE is the change in import competition exposure.

Table A.2: Brazilian workforce during 1990s trade liberalization

	Population Census Year	
	1991	2000
N	51,332,184 (42.6%)	69,295,119 (57.4%)
Formal wage work	0.403 (0.491)	0.312 (0.463)
Informal wage work	0.324 (0.468)	0.344 (0.475)
Self-employed	0.239 (0.427)	0.205 (0.404)
Unemployed	0.044 (0.206)	0.138 (0.345)
Hours worked	43.534 (12.007)	43.950 (14.655)
Earnings (2000 reais)	331.088 (412.028)	553.742 (704.550)
Earnings (2000 reais) - Formal	383.827 (413.412)	596.306 (653.343)
Earnings (2000 reais) - Informal	279.714 (418.674)	497.428 (725.555)
Earnings (2000 reais) - Self-employed	309.382 (388.172)	568.641 (743.992)
Minimum wage (2000 reais)	126.514 (0.000)	151.000 (0.000)
Female	0.326 (0.469)	0.404 (0.491)
Age	34.262 (11.490)	34.717 (11.500)
Years of education	5.899 (4.522)	7.032 (4.392)

Notes: Descriptive statistics from the 1991 and 2000 population censuses, computed from the sample extracts in the supplementary materials of Dix-Carneiro and Kovak (2017) with individual sample weights. The sample is workers employed in the private sector. Formal wage work is having a signed *Carteira de Trabalho* (variables v350 and v0447 in the 1991 and 2000 censuses). Earnings and hours are for the main job; earnings statistics exclude the top and bottom 1% of the earnings distribution and are in 2000 reais (IPCA deflator). The minimum wage is the federal monthly minimum wage for July 1991 and for 2000, for a 44-hour work week.

Table A.3: Formal employment and worker flows by year

	1986	1987	1988	1989	1990	1991	1992	1993	1994	1995	1996	1997	1998	1999	2000
Panel A: December 31 Employment (millions)															
Total	10.64	20.80	21.04	21.74	20.79	19.65	18.25	18.84	20.37	23.54	23.29	24.09	23.93	23.64	25.25
Exporters	1.69	3.97	3.88	4.17	3.88	3.59	3.27	3.35	3.47	3.53	3.06	2.79	2.46	2.21	2.26
Non-Exp. Tradables	2.18	3.67	3.49	3.64	3.52	3.25	2.91	3.14	3.96	4.56	4.56	4.85	4.84	4.94	5.40
Non-Tradables	6.77	13.16	13.66	13.94	13.39	12.81	12.07	12.35	12.94	15.45	15.67	16.45	16.63	16.49	17.59
Panel B: Flows In/Out of Formal Sector (millions)															
Hired from Outside	7.84	3.60	3.73	4.01	3.47	3.23	2.95	3.47	4.46	4.63	4.04	4.53	4.37	4.33	5.03
Into Exporters	1.20	0.50	0.45	0.55	0.40	0.39	0.32	0.40	0.44	0.41	0.29	0.27	0.21	0.22	0.25
Into Non-Exp. Tradables	1.68	0.70	0.68	0.80	0.69	0.61	0.55	0.70	1.03	1.04	0.90	1.03	0.98	1.01	1.18
Into Non-Tradables	4.97	2.40	2.60	2.66	2.38	2.23	2.08	2.37	2.98	3.17	2.85	3.23	3.18	3.10	3.60
Separated to Outside	3.00	3.17	3.17	3.43	4.11	3.92	3.40	2.86	3.16	4.19	3.79	3.89	4.13	3.66	3.71
From Exporters	0.40	0.48	0.41	0.48	0.64	0.57	0.46	0.38	0.41	0.54	0.40	0.38	0.34	0.23	0.21
From Non-Exp. Tradables	0.59	0.66	0.59	0.64	0.78	0.74	0.62	0.53	0.67	0.92	0.82	0.84	0.90	0.81	0.85
From Non-Tradables	2.01	2.04	2.17	2.31	2.69	2.61	2.33	1.94	2.08	2.72	2.57	2.67	2.89	2.62	2.65
Panel C: Flows Between Firms (millions)															
From Exporters	0.88	0.96	0.79	0.82	0.83	0.63	0.51	0.52	0.65	0.74	0.53	0.49	0.41	0.36	0.35
From Non-Exp. Tradables	1.08	1.11	0.97	0.94	0.90	0.71	0.58	0.65	0.94	1.18	1.05	1.08	1.02	1.05	1.19
From Non-Tradables	3.83	3.73	3.75	3.55	3.27	2.80	2.36	2.59	3.19	3.92	3.57	3.58	3.46	3.55	3.92

Notes: This table summarizes formal private sector employment and worker flows in Brazil from 1986 to 2000 using matched employer-employee data from RAIS. Panel A reports December 31 employment (in millions) by firm trading status: exporters, non-exporting tradable firms, and non-tradable firms. Panel B reports annual flows into and out of the formal sector, separately by destination or origin firm type. "Hired from Outside" counts workers appearing in the formal sector on December 31 who were not employed in any formal firm on December 31 of the previous year. "Separated to Outside" counts workers employed in a formal firm on January 1 who are not found in any formal firm on December 31 of the same year. Panel C reports within-year job-to-job transitions between formal sector firms, by origin firm type. All values are in millions of workers.

Table A.4: Effect of import competition on market outcomes: Alternative specifications

	Main specification	ICE weights are firms' 1991 linear payroll shares	Dropping the largest 1% of markets	Unweighted
	(1)	(2)	(3)	(4)
Panel A: Formal sector labor market concentration				
Δ Payroll Herfindahl (premium-based)	0.172	0.421	0.174	0.212
	(0.042)	(0.074)	(0.031)	(0.017)
Δ Payroll Herfindahl	0.183	0.439	0.179	0.213
	(0.045)	(0.083)	(0.031)	(0.017)
Δ Employment Herfindahl	0.109	0.278	0.134	0.239
	(0.024)	(0.035)	(0.022)	(0.016)
Panel B: Number of firms and formal employment				
Δ Log number of firms	−0.971	−1.443	−0.424	−0.510
	(0.188)	(0.196)	(0.099)	(0.048)
Δ Log formal employment	−0.398	−0.048	0.194	−0.401
	(0.186)	(0.199)	(0.104)	(0.069)
Panel C: Wages				
Δ Log average firm wage premium	−0.124	0.200	0.112	−0.071
	(0.130)	(0.151)	(0.060)	(0.035)
Δ Log formal wage premium	−0.105	0.208	0.124	0.027
	(0.089)	(0.113)	(0.054)	(0.030)
Δ Log formal average wage	−0.075	0.255	0.119	0.014
	(0.110)	(0.131)	(0.062)	(0.035)
Weighted by 1991 formal employment	Yes	Yes	Yes	No
Observations	296,311	296,311	293,341	296,311
Local labor markets	19,757	19,757	19,559	19,757

Notes: Standard errors in parentheses, clustered at the local labor market level. Each cell reports the 2000 coefficient of the stacked differences-in-differences regression of Table 1, with local labor markets weighted by their 1991 formal employment except in column (4). Column (1) repeats Table 1, column (1). Column (2) constructs ΔICE_m with firms' 1991 linear payroll shares as weights, instead of squared payroll shares. Column (3) drops the largest 1% of local labor markets by 1991 formal employment. Column (4) is the main specification without weights, each local labor market counting once; its effect on concentration is identified mostly by comparing markets with and without tradable firms (Appendix D) and is the one that enters the effect on markdowns of Table 4. The sample is that of the concentration regressions.

Table A.5: Effects of import competition on market outcomes: Alternative fixed effects and controls

	Year FE (main) (1)	State × year FE (2)	State × year FE and 1986–1990 trends (3)	Microregion × year FE (4)
Panel A: Formal sector labor market concentration				
Δ Payroll Herfindahl (premium-based)	0.172	0.146	0.152	0.084
	(0.042)	(0.036)	(0.036)	(0.035)
Δ Payroll Herfindahl	0.183	0.152	0.157	0.092
	(0.045)	(0.038)	(0.038)	(0.038)
Δ Employment Herfindahl	0.109	0.107	0.110	0.044
	(0.024)	(0.023)	(0.023)	(0.021)
Panel B: Number of firms and formal employment				
Δ Log number of firms	−0.971	−0.795	−0.681	−0.440
	(0.188)	(0.138)	(0.129)	(0.122)
Δ Log formal employment	−0.398	−0.200	−0.194	−0.027
	(0.186)	(0.146)	(0.145)	(0.150)
Panel C: Wages				
Δ Log average firm wage premium	−0.124	−0.198	−0.190	−0.248
	(0.130)	(0.115)	(0.118)	(0.136)
Δ Log formal wage premium	−0.105	−0.096	−0.084	−0.085
	(0.089)	(0.076)	(0.077)	(0.084)
Δ Log formal average wage	−0.075	−0.154	−0.154	−0.196
	(0.110)	(0.096)	(0.098)	(0.108)

Notes: Standard errors in parentheses, clustered at the local labor market level. Each entry is the coefficient on ΔICE_m in the final year of the stacked differences-in-differences regression of equation (12). Column (1) is the main specification, with local labor market and year fixed effects. Column (2) replaces year fixed effects with state × year fixed effects. Column (3) adds the market's 1986–1990 change in the outcome, interacted with year. Column (4) uses microregion × year fixed effects. Markets are weighted by their 1991 formal employment, as in Table 1, column (1).

Table A.6: Effect of import competition on market outcomes: Alternative clustering

	Main specification	Two-way clustered by microregion and occupational group	AKM (2019) standard errors
	(1)	(2)	(3)
Panel A: Formal sector labor market concentration			
Δ Payroll Herfindahl (premium-based)	0.172	0.172	0.172
	(0.042)	(0.046)	(0.045)
Δ Payroll Herfindahl	0.183	0.183	0.183
	(0.045)	(0.048)	(0.048)
Δ Employment Herfindahl	0.109	0.109	0.109
	(0.024)	(0.035)	(0.033)
Panel B: Number of firms and formal employment			
Δ Log number of firms	−0.971	−0.971	−0.971
	(0.188)	(0.576)	(0.407)
Δ Log formal employment	−0.398	−0.398	−0.398
	(0.186)	(0.442)	(0.371)
Panel C: Wages			
Δ Log average firm wage premium	−0.124	−0.124	−0.124
	(0.130)	(0.086)	(0.122)
Δ Log formal wage premium	−0.105	−0.105	−0.105
	(0.089)	(0.099)	(0.120)
Δ Log formal average wage	−0.075	−0.075	−0.075
	(0.110)	(0.070)	(0.124)
Observations	296,311	296,311	296,311
Local labor markets	19,757	19,757	19,757

Notes: Standard errors in parentheses. Point estimates are identical across columns; only standard errors differ. Column 1 clusters at the local labor market level, column 2 two-way clusters by microregion and occupational group, and column 3 uses AKM (2019) shift-share standard errors. All columns report the 2000 coefficient from stacked differences-in-differences regressions, as in Table 1.

Table A.7: Effect of import competition on concentration: Sources

Panel A: Exit, firms present in both years, and entry	
Δ Payroll Herfindahl	0.183
	(0.045)
Via firms present in 1991 only (exit)	−0.066
	(0.035)
Via the shares of firms present in both years	0.127
	(0.053)
Via firms present in 2000 only (entry)	0.122
	(0.040)
1991 payroll share of firms absent in 2000	−0.020
	(0.070)
2000 payroll share of firms absent in 1991	−0.253
	(0.077)
Observations	296,370
Panel B: By firm type	
Δ Payroll Herfindahl	0.183
	(0.045)
Exporters	0.144
	(0.041)
Firms present in 1991 only (exit)	0.060
	(0.018)
Firms present in both years	0.095
	(0.038)
Firms present in 2000 only (entry)	−0.012
	(0.004)
Non-exporting tradable firms	−0.005
	(0.012)
Firms present in 1991 only (exit)	0.007
	(0.005)
Firms present in both years	0.012
	(0.003)
Firms present in 2000 only (entry)	−0.024
	(0.012)
Non-tradable firms	0.044
	(0.019)
Firms present in 1991 only (exit)	0.084
	(0.016)
Firms present in both years	0.020
	(0.011)
Firms present in 2000 only (entry)	−0.060
	(0.011)
Observations	296,370

Notes: Each cell is the 2000 coefficient on $\Delta ICE_m \times \mathbf{1}[t = 2000]$ from the main specification of Table 1, column (1) (equation 12, 1986–2000, market and year fixed effects, markets weighted by their 1991 formal employment, standard errors clustered by market in parentheses), with the outcome in each row. Panel A: the first row is the change in the market's December wage-bill Herfindahl since 1991, the Payroll Herfindahl of Table 1; the next three rows decompose it exactly, firm by firm, into the change in the Herfindahl of the 1991 shares when the firms absent in t are removed and the remaining shares renormalized (exit), the change between the renormalized 1991 shares and the renormalized t shares of the firms present in both years, and the change from adding the firms present in t only (entry). The last two rows are the 1991 payroll share of the firms absent in t and the payroll share in t of the firms absent in 1991. Panel B: the change in the Herfindahl is the sum over firms of the change in the firm's squared payroll share, with a share of zero in a year in which the firm is absent; each row sums these changes over the firms of one type (exporters, firms that exported at any point in 1991–1994; other tradable firms; non-tradable firms) and, within type, over the firms absent in 2000, present in both years, and absent in 1991. Shares are not renormalized in Panel B, so the rise in a surviving firm's share when other firms leave is in the row of the surviving firm. A 10% increase in exposure is 0.1 units of ΔICE_m.

Table A.8: Within-market inverse elasticity $1/\eta$: Robustness

	Main specification (all-worker premia)	Stayers' wage premia	December average wage	Including exiting firms, coding employment and wages at exit as zero
	(1)	(2)	(3)	(4)
Panel A: First stage				
Δ Firm log employment in LLM	-0.736	-0.743	-0.736	-0.737
	(0.076)	(0.091)	(0.076)	(0.075)
First stage F	94.915	66.474	94.915	96.505
Panel B: Reduced form				
Δ Firm wage premium in LLM	-0.612	-0.579	-0.604	-0.617
	(0.036)	(0.049)	(0.038)	(0.036)
Panel C: 2SLS				
Labor supply within-market cross-firm	0.832	0.779	0.820	0.837
inverse elasticity of substitution	(0.093)	(0.107)	(0.094)	(0.093)
Implied upper bound on wage take-home share	55%	56%	55%	54%
Local labor market (LLM) FE	Yes	Yes	Yes	Yes
Pre-trend controls (1986–91)	Yes	Yes	Yes	Yes
Observations	363,407	190,069	363,407	700,065
Firms	149,083	89,962	149,083	303,393
Local labor markets	11,632	9,502	11,632	14,175

Notes: Standard errors clustered by firm in parentheses. All regressions include local labor market (LLM) fixed effects and are weighted by baseline market employment. Column (1): main specification (Table 2, column (1)), using the December wage premium of all of the firm's workers, conditional on observables. Column (2): wage premium estimated on the firm's stayers, workers in the same firm, microregion, and occupation in 1991 and 2000; Table 2, column (2), reports the version conditional on worker fixed effects. Column (3): December average wage instead of the residualized wage premium. Column (4): includes firms that exited between 1991 and 2000, coding their exit-year employment and wages as zero (inverse hyperbolic sine transform). Table 2 reports microregions as local labor markets and Appendix Table E.2 the effective rate of protection as the shock.

Table A.9: Alternative standard errors for the within-market and cross-market inverse elasticities of substitution

	$1/\eta$ (Table 2, column (1))			$1/\theta$ (Table 3, column (1))	
	Clustered by firm (main)	Clustered by local labor market	Clustered by sector (CNAE)	Clustered by local labor market (main)	Two-way clustered by microregion and occupation
	(1)	(2)	(3)	(4)	(5)
Panel A: First stage					
Δ Firm log employment in LLM	-0.736	-0.736	-0.736		
	(0.076)	(0.088)	(0.148)		
Δ LLM employment index				-0.294	-0.294
				(0.044)	(0.100)
First stage F	94.9	70.0	24.8	45.6	8.7
Panel B: Reduced form					
Δ Firm wage premium in LLM	-0.612	-0.612	-0.612		
	(0.036)	(0.114)	(0.087)		
Δ LLM wage premium index				-0.183	-0.183
				(0.050)	(0.103)
Panel C: 2SLS					
$1/\eta$	0.832	0.832	0.832		
	(0.093)	(0.218)	(0.132)		
$1/\theta - 1/\eta$				0.622	0.622
				(0.217)	(0.473)
$1/\theta$				1.454	1.454
				(0.196)	(0.464)
Observations	363,407	363,407	363,407	17,628	17,628

Notes: Standard errors in parentheses. Point estimates are those of Table 2, column (1) (columns (1)–(3)) and of Table 3, column (1) (columns (4)–(5)); only the clustering of the standard errors changes. Column (1) clusters by firm, column (2) by local labor market (microregion × occupation), column (3) by sector (CNAE95); column (4) clusters by local labor market, column (5) two-way by microregion and occupational group.

Table A.10: Correlates of firm-level wage markdowns

	Dependent variable: $\ln(\mu_{zm})$			
	(1)	(2)	(3)	(4)
Panel A: Employment				
$\ln(\text{Employment}_{zm})$	0.003824	0.003599	0.003423	0.003882
	(0.000185)	(0.000171)	(0.000153)	(0.000173)
Observations	9,484,956	9,484,955	9,380,001	9,484,955
Panel B: Residual wage				
$\ln(\text{Residual wage}_{zm})$	0.000759	0.000666	0.000313	0.003355
	(0.000152)	(0.000121)	(0.000111)	(0.000156)
Observations	9,484,956	9,484,955	9,380,001	9,484,955
Panel C: Average wage				
$\ln(\text{Average wage}_{zm})$	0.000710	0.000553	0.000030	0.002825
	(0.000141)	(0.000113)	(0.000098)	(0.000142)
Observations	9,484,956	9,484,955	9,380,001	9,484,955
Year FE	Yes	Yes	Yes	Yes
Sector FE		Yes		
Firm FE			Yes	
Local labor market FE				Yes
Firms	580,576	580,575	475,621	580,576
Local labor markets	19,023	19,023	19,005	19,022

Notes: This table presents regressions of log firm-level wage markdowns on firm-level employment, residual wages, and average wages. All regressions include year fixed effects. Standard errors in parentheses, clustered by local labor market.

Table A.11: Heterogeneity by informality: quartile 1 split at its median and quartiles 2–4 pooled

	All markets	Quartile 1 (most formal)			Quartiles 2–4 (pooled)
		All	Upper half	Lower half	
	(1)	(2)	(3)	(4)	(5)
Share of national formal employment, 1991	0.995	0.726	0.573	0.153	0.269
Median formal share of employment, 1991	0.265	0.502	0.559	0.457	0.205
Panel A: Effect of 10% increase in import competition on concentration					
Δ Payroll Herfindahl (premium-based)	0.017	0.015	0.013	0.018	0.019
	(0.004)	(0.005)	(0.006)	(0.009)	(0.004)
Mean (unweighted), 1991	0.287	0.212	–	–	0.331
Mean (weighted), 1991	0.079	0.064	0.053	0.118	0.151
Observations	296,311	108,396	57,628	50,768	187,915
Local labor markets	19,757	7,227	3,842	3,385	12,530
Panel B: Within-market cross-firm inverse elasticity $1/\eta$ (2SLS)					
$1/\eta$	0.832	0.824	0.821	1.650	-1.252
	(0.093)	(0.091)	(0.091)	(1.401)	(0.977)
First-stage F	94.9	97.1	97.0	1.4	1.9
Implied upper bound on take-home share	0.546	0.548	0.549	–	–
Observations	363,407	287,985	237,348	50,637	75,422
Firms	149,083	115,676	95,720	20,808	34,513
Local labor markets	11,632	5,361	3,054	2,307	6,271

Notes: See notes to Table 6. Columns (3) and (4) split quartile 1 at the median formal share of total employment across its microregions. Column (5) pools quartiles 2 to 4.

Table A.12: Wage effects at firms facing larger tariff cuts vs the market average

	Δ Log wage premium	Δ Log employment	Δ Payroll share of the market	Δ Log wage take-home share	Δ Log marginal revenue product of labor
	(1)	(2)	(3)	(4)	(5)
Panel A. Firms facing larger tariff cuts, within local labor markets					
Δ in log import tariff faced by firm	-0.612	-0.736	-0.0004	0.0001	-0.613
	(0.036)	(0.076)	(0.0001)	(0.0000)	(0.036)
Share of the wage premium response	100%	–	–	0%	100%
Local labor market FE, pre-trend controls	Yes	Yes	Yes	Yes	Yes
Observations (firm × market)			363,407		
Panel B. Local labor market average, across local labor markets					
ΔICE_m	-0.124	-0.398	–	–	–
	(0.130)	(0.186)			
Local labor market and year FE	Yes	Yes			
Observations (market × year)			296,311		

Notes: Standard errors in parentheses, clustered by firm in Panel A and by local labor market in Panel B. All regressions are weighted by the market's 1991 formal employment. In Panel A, each column reports the coefficient on the firm's 1990–1994 change in log import tariff in a regression of the firm's 1991–2000 change in the column's outcome, with local labor market fixed effects and the pair's 1986–1991 changes in log employment and in the log wage premium, on the sample of Table 2, column (1). Column (1) is the reduced form of Table 2 for the firm's wage premium, and column (2) its first stage. Column (3) is the change in the firm's share of its market's payroll. Column (4) is the change in the firm's log wage take-home share implied by its payroll share and the estimated elasticities, $\ln \mu_{zm}^{-1} = -\ln(1 + (1/\eta)(1 - s_{zm}) + (1/\theta)s_{zm})$, with $1/\eta$ from Table 2, column (1) and $1/\theta$ from Table 3. Column (5) is column (1) minus column (4), the change in the log marginal revenue product of labor. The row "share of the wage premium response" divides columns (4) and (5) by column (1). Panel B reports the estimates of Table 1, column (1): the 2000 coefficient on ΔICE_m for the market's log average firm wage premium (column 1) and log employment (column 2), relative to 1991.

Table A.13: Extended model: Markdown effects by formality quartile

	All microregions	Quartile 1 (most formal)	Quartile 2	Quartile 3	Quartile 4 (least formal)
	(1)	(2)	(3)	(4)	(5)
Microregions	475	121	121	116	117
Share of national formal employment, 1991	0.997	0.726	0.171	0.073	0.027
Share of 1991 formal employment, RAIS (weight)	1.000	0.834	0.119	0.039	0.008
Panel A: Market structure, 1991 (weighted averages)					
Expected payroll Herfindahl	0.022	0.015	0.048	0.091	0.121
Wage work share of market payroll	0.780	0.800	0.707	0.638	0.537
Panel B: Wage take-home share at the 1991 averages					
Extended model	0.531	0.531	0.530	0.529	0.530
with all market payroll in wage work ($s_m = 1$)	0.530	0.531	0.528	0.524	0.522
with $\Omega = 0$ ($1/\eta$ in place of $1/\tilde{\eta}$)	0.533	0.534	0.532	0.531	0.532
Panel C: Effect of import competition on the average wage take-home share, at the 1991 averages					
Δ average inverse elasticity, per 10% increase in the trade shock	0.0053	0.0057	0.0038	0.0021	0.0002
	(0.0014)	(0.0015)	(0.0012)	(0.0009)	(0.0006)
through the change in concentration	0.0056	0.0059	0.0046	0.0036	0.0021
through the change in the wage work share	-0.0004	-0.0002	-0.0008	-0.0015	-0.0020
Implied change in take-home share (pp)	-0.15	-0.16	-0.11	-0.06	-0.00

Notes: Microregions are grouped into quartiles of the 1991 formal share of total employment, as in Table 6. The take-home share is equation 24 evaluated at the group's 1991 averages of the expected payroll Herfindahl and of the wage work share of market payroll, with the elasticities of Table 7; the two counterfactual rows set the wage work share to one and the separation-risk bias Ω to zero. The effect of a 10% increase in the trade shock is equation 25 evaluated at the same averages, with the effects on the expected Herfindahl and on the wage work share from the microregion regressions of Table 7, Panels A–B, weighted by 1991 formal employment and estimated separately by group; the two rows below it are its concentration and wage-work-share components. Standard errors by the delta method, holding the elasticities fixed.

Table A.14: Extended model: Distribution of the wage take-home share and of the effect of trade across microregions

	Min	p5	p10	p25	p50	p75	p90	p95	Max
Panel A: 1991									
Wage take-home share	0.492	0.521	0.525	0.527	0.530	0.531	0.532	0.534	0.562
Expected payroll Herfindahl	0.002	0.012	0.017	0.031	0.070	0.145	0.292	0.410	0.930
Wage work share of market payroll	0.070	0.324	0.437	0.587	0.694	0.771	0.806	0.823	0.874
N (microregions)					475				
Panel B: 2000									
Wage take-home share	0.494	0.519	0.525	0.528	0.531	0.532	0.532	0.532	0.532
Expected payroll Herfindahl	0.001	0.006	0.008	0.014	0.032	0.081	0.173	0.277	1.000
Wage work share of market payroll	0.475	0.600	0.637	0.686	0.731	0.766	0.787	0.801	0.894
N (microregions)					475				
Panel C: Effect of a 10% increase in the trade shock (1991 market structure)									
Change in take-home share (pp)	-0.18	-0.16	-0.15	-0.13	-0.09	-0.01	0.09	0.17	0.48
N (microregions)					475				

Notes: Unweighted distribution across the 475 microregions of the extended model. The take-home share is equation 24 evaluated at each microregion's expected payroll Herfindahl and wage work share of market payroll in the year, with the elasticities of Table 7, which are common to all microregions. Panel C is equation 25 evaluated at each microregion's 1991 market structure, with the effects on the expected Herfindahl and on the wage work share of Table 7, Panels A and B.

A Data and Methods Appendix

Data on workers and firms: RAIS

Overview. I use Brazil's Relação Anual de Informações Sociais (RAIS) for years 1986 to 2000 as my source of information on workers and firms. RAIS is an administrative employer-employee linked dataset collected by the federal government for the purposes of administering workers' social security. Thus, RAIS covers all workers with signed worker cards (Carteira do Trabalho), namely the entirety of formal sector employment. Firms report to RAIS once a year, listing all workers who ever worked for the firm in the prior calendar year. Firms are required to report a rich set of information about each employment contract (e.g., occupation, admission date, separation date, etc.), as well as worker demographics (i.e., education, date of birth, and gender), separately for each establishment. The municipality of each establishment as well as the economic sector of the firm are also reported.

Wages. RAIS includes two wage variables for years 1986–2000: average monthly earnings and December monthly earnings. Both variables are reported as multiples of the national minimum wage.

Occupation codes. RAIS' occupation codes are 5-digit variables "CBO" (prior to 1994) and "CBO94" (1994 onwards). I focus on the first 2 digits to group workers into occupation groups. Both variables share the same data dictionary, with the only difference between them being phased-out and phased-in occupation codes. I have compiled a complete list of all raw occupation codes, along with the total number of workers in each of them, labels, and flags for which codes were either "phased-out" or phased-in, which I identified based on whether the number of workers changed by a factor of more than 100 between any two years. I then re-classified the first two digits of all phased-out and phased-in codes as "99 - Other occupations," a reclassification that affects roughly 10% of all workers.

Sector codes. RAIS' finest sector codes for 1986–2000 are 4-digit "IBGESUBATIVIDADE" (prior to 1995) and 5-digit "CNAE95" (1995 onwards). I focus on the 5-digit CNAE95 codes to map tariff shocks to firms in RAIS. For firms that exit the data prior to reporting any CNAE95 codes, I assign a CNAE95 code using a correspondence table I constructed using the pre-1995 and post-1995 codes of firms in business in both periods. To each IBGESUBATIVIDADE code I assign the most commonly reported CNAE95 code. Finally, throughout all years I use the first CNAE95 code ever reported by a firm as its official CNAE95 code.

Sample restrictions. I focus on workers employed as of December 31 of each year, aged 18–64, and with positive December earnings. I exclude all workers in the public sector or with unknown sector. To make sure all public sector workers are excluded, I further exclude workers whose employer's economic activity was not marked as government, but who hold public sector occupations (i.e., Diplomats, Civil servants, and Post office). Finally, following Dix-Carneiro and

Kovak (2017) I exclude from all analyses the microregion of Manaus.[56]

Data on tariff shocks: TRAINS

I use tariff data from UNCTAD's Trade Analysis Information System (TRAINS), which I download from the World Integrated Trade Solution (WITS) website (https://wits.worldbank.org/). The raw tariff data are available for Brazil at the 8-digit HS product level for years 1988 (the first year the data are available) through 2000. As outlined in Section 4, I compute a firm's tariff reduction shock as the change in log one plus the firm's CNAE95 sector code's nominal tariff between years 1990 and 1994.

To map product-level tariffs to CNAE95, I use Brazil's NCM (Nomenclatura Comum do Mercosul) classification, whose first six digits correspond to the HS. I map NCM codes to CNAE95 using official IBGE correspondence tables.[57] Over 90% of products map to a single activity code; I randomly assign the remainder across the handful of activity codes to which they correspond per the concordance tables. I compute CNAE95-level tariffs as averages of product-level nominal tariffs, weighted by the number of tariff lines. For robustness, I compute effective rates of protection (ERP) using the Corden formula: $ERP_j = (t_j - \sum_i a_{ij} t_i)/(1 - \sum_i a_{ij})$, where t_j is sector j's nominal tariff, t_i are input tariffs, and a_{ij} are Leontief coefficients from Brazil's 1985 intersectoral coefficients matrix ("Tabela 20") at Nível 50.[58] I map Nível 50 sectors to CNAE95 using CONCLA correspondence tables.

Other data

List of exporters. I classify firms as exporters during the reform period (1990–1994) by matching the list of exporters during that period to RAIS using firms' unique identifiers (CNPJ). The list of exporters was provided by the (extinct as of 2019) Ministry of Development, Industry, and Foreign Trade, currently a part of the Ministry of the Economy, in October 2018.

Census. I use data from the 1991 and 2000 census, available in the supplementary materials from Dix-Carneiro and Kovak (2017).

[56] Law 8.387 of December 1991 changed the rules of the Manaus Free Trade Zone during the liberalization, introducing a minimum set of production operations as a condition for its tax benefits and an 88 percent reduction of the import duty on approved projects' foreign inputs, so changes in Manaus over the period combine the liberalization with a change in the zone's own regime. Using synthetic controls, Possebom (2017) finds that the zone's effects on the region's GDP per capita were largest between 1975 and 1990 and weakened after 1991, as national tariff cuts reduced its relative advantage.

[57] https://concla.ibge.gov.br/classificacoes/correspondencias/atividades-economicas.

[58] https://www.ibge.gov.br/estatisticas-novoportal/economicas/contas-nacionais/9085-matriz-de-insumo-produto.html?&t=downloads.

Methods: wage premia regressions

For each year, I estimate each firm's wage premium in its local labor market as firm × market fixed effects in a regression of worker log December earnings on the firm × market fixed effects and the same worker observable controls as Dix-Carneiro and Kovak (2017), namely: a dummy for female; 4 age group dummies (25–29; 30–39; 40–49; 50–64); 8 education group dummies (primary school, incomplete primary school, middle school, incomplete middle school, high school, incomplete high school, college, incomplete college). The omitted category is therefore males aged 18–24 with no formal education. Similarly, for each year, I estimate each market's wage premia from a regression of worker log December earnings on the market fixed effects and the previously mentioned worker observable controls.

For my main estimates of the within-market cross-firm elasticity (column (1) of Table 2), I use as firm wage premia the firm × market fixed effects of these regressions, estimated on all of the firm's workers, and my outcome is their 1991–2000 change. As robustness, I instead estimate the premia on the firm's stayers—workers employed in the same firm, microregion, and occupation in both 1991 and 2000—by regressing stayers' log December earnings on firm × market fixed effects and the worker observable controls above for each of the two years (Appendix Table A.8); in column (2) of Table 2, I additionally condition on worker fixed effects. In the extended model of Section 8, I use the analogous premia for firm × microregion pairs.

For the robustness exercise of the within-market cross-firm elasticity using wage premia that condition on worker fixed effects (Table 2, column (2)), I estimate each firm's wage premia in 1991 and 2000 as firm × market × year fixed effects in a regression—containing years 1991 and 2000, on stayers—of worker log December earnings on worker fixed effects, the firm × market × year fixed effects, and worker observable-characteristics-by-year controls.

B Model Appendix

This Appendix provides detailed derivations for various expressions in Section 2.

B.1 Derivation of labor supply equation

Consider an economy consisting of a continuum of homogeneous workers j, a large but finite number of local labor markets m, and a finite number of firms z within each local labor market. Each worker chooses to which firm-market pair zm they provide h^j_{zm} units of labor by minimizing their indirect disutility of work V_{zm} subject to making reservation earnings $y^j \sim F(y)$:

$$\begin{aligned}\min_{zm} V^j_{zm} &= \ln l^j_{zm} + \ln \xi_m + \ln \xi_{zm} - \xi^j_{zm} \\ &\text{s.t. } l^j_{zm} w_{zm} \geq y_j\end{aligned}$$

where ξ^j_{zm} is an idiosyncratic taste for working at firm z in market m, and ξ_m and ξ_{zm} are taste shifters common to all workers. This is equivalent to

$$\max_{zm} \ln w_{zm} - \ln y_j - \ln \xi_m - \ln \xi_{zm} + \xi^j_{zm}$$

Now suppose ξ^j_{zm} has the following Gumbel distribution, a member of the General Extreme Value (GEV) family:

$$G\left(\left\{\xi^j_{zm}\right\}\right) = \exp\left[-\sum_m \left(\sum_{z\in B_m} e^{-\left(1+\frac{\sigma}{1-\sigma}\right)\xi^j_{zm}}\right)^{\frac{1+\frac{\varphi}{1-\varphi}}{1+\frac{\sigma}{1-\sigma}}}\right]$$

where $0 \leq \sigma < 1$ is the index of similarity across firms within a market, $0 \leq \varphi < 1$ is the index of similarity across markets, and B_m is the set of firms in market m.

The probability that worker j chooses firm z in market m is

$$P\left(\xi^j_{zm} > \ln w_{zm} - \ln y_j - \ln \xi_m - \ln \xi_{zm}\right),$$

which can be decomposed as:

$$P^j_{zm} = P(z|B_m)\, P(B_m) \quad \forall j$$

where $P(z|B_m)$ is the probability of choosing firm z conditional on choosing market m with set B_m of firms, and $P(B_m)$ is the probability of choosing market m. By the results in McFadden

(1978), P_{zm} can be computed as:

$$\begin{aligned}
P\left(z|B_m\right) =& \frac{\exp\left[\left(\ln w_{zm} - \ln y^j - \ln \xi_m - \ln \xi_{zm}\right)/(1-\sigma)\right]}{\sum_{k\in B_n}\exp\left[\left(\ln w_{kn} - \ln y^j - \ln \xi_m - \ln \xi_{km}\right)/(1-\sigma)\right]} \\
=& \frac{\exp\left[\left(\ln w_{zm}^{\frac{1}{1-\sigma}} - \ln y^{j\,\frac{1}{1-\sigma}} - \ln \xi_m^{\frac{1}{1-\sigma}} - \ln \xi_{zm}^{\frac{1}{1-\sigma}}\right)\right]}{\sum_{k\in B_m}\exp\left[\left(\ln w_{km}^{\frac{1}{1-\sigma}} - \ln y^{j\,\frac{1}{1-\sigma}} - \ln \xi_m^{\frac{1}{1-\sigma}} - \ln \xi_{km}^{\frac{1}{1-\sigma}}\right)\right]} \\
=& \frac{\left(\frac{w_{zm}}{y^j \xi_m \xi_{zm}}\right)^{\frac{1}{1-\sigma}}}{\sum_{k\in B_n}\left(\frac{w_{km}}{y^j \xi_m \xi_{km}}\right)^{\frac{1}{1-\sigma}}} \\
=& \frac{\left(\frac{w_{zm}}{\xi_{zm}}\right)^{\frac{1}{1-\sigma}}}{\sum_{k\in B_n}\left(\frac{w_{km}}{\xi_{km}}\right)^{\frac{1}{1-\sigma}}}
\end{aligned}$$

and

$$\begin{aligned}
P\left(B_m\right) =& \frac{\left\{\sum_{z\in B_m}\exp\left[\left(\ln w_{zm} - \ln y^j - \ln \xi_m - \ln \xi_{zm}\right)/(1-\sigma)\right]\right\}^{\frac{1-\sigma}{1-\varphi}}}{\sum_l\left\{\sum_{k\in B_l}\exp\left[\left(\ln w_{kl} - \ln y^j - \ln \xi_l - \ln \xi_{kl}\right)/(1-\sigma)\right]\right\}^{\frac{1-\sigma}{1-\varphi}}} \\
=& \frac{\left[\sum_{z\in B_m}\left(\frac{w_{zm}}{y^i \xi_m \xi_{zm}}\right)^{\frac{1}{1-\sigma}}\right]^{\frac{1-\sigma}{1-\varphi}}}{\sum_l\left[\sum_{k\in B_l}\left(\frac{w_{km}}{y^i \xi_m \xi_{km}}\right)^{\frac{1}{1-\sigma}}\right]^{\frac{1-\sigma}{1-\varphi}}} \\
=& \frac{\left[\left(\frac{1}{\xi_m}\right)^{\frac{1}{1-\sigma}}\sum_{z\in B_m}\left(\frac{w_{zm}}{\xi_{zm}}\right)^{\frac{1}{1-\sigma}}\right]^{\frac{1-\sigma}{1-\varphi}}}{\sum_l\left[\left(\frac{1}{\xi_l}\right)^{\frac{1}{1-\sigma}}\sum_{k\in B_l}\left(\frac{w_{kl}}{\xi_{kl}}\right)^{\frac{1}{1-\sigma}}\right]^{\frac{1-\sigma}{1-\varphi}}}
\end{aligned}$$

Putting them together

$$P_{zm}^j = \frac{\left(\frac{w_{zm}}{\xi_{zm}}\right)^{\frac{1}{1-\sigma}}}{\sum_{k\in B_n}\left(\frac{w_{km}}{\xi_{km}}\right)^{\frac{1}{1-\sigma}}} \times \frac{\left[\left(\frac{1}{\xi_m}\right)^{\frac{1}{1-\sigma}}\sum_{z\in B_m}\left(\frac{w_{zm}}{\xi_{zm}}\right)^{\frac{1}{1-\sigma}}\right]^{\frac{1-\sigma}{1-\varphi}}}{\sum_l\left[\left(\frac{1}{\xi_l}\right)^{\frac{1}{1-\sigma}}\sum_{k\in B_l}\left(\frac{w_{kl}}{\xi_{kl}}\right)^{\frac{1}{1-\sigma}}\right]^{\frac{1-\sigma}{1-\varphi}}} \quad \forall j$$

Let $\eta \equiv \frac{\sigma}{1-\sigma} > 0$, $\theta \equiv \frac{\varphi}{1-\varphi} > 0$, and denote $P^j_{zm} = P_{zm}$ for simplicity. Then:

$$P_{zm} = \frac{\left(\frac{w_{zm}}{\xi_{zm}}\right)^{1+\eta}}{\sum_{k\in B_n}\left(\frac{w_{km}}{\xi_{km}}\right)^{1+\eta}} \times \frac{\left[\left(\frac{1}{\xi_m}\right)^{1+\eta}\sum_{z\in B_m}\left(\frac{w_{zm}}{\xi_{zm}}\right)^{1+\eta}\right]^{\frac{1+\theta}{1+\eta}}}{\sum_l\left[\left(\frac{1}{\xi_l}\right)^{1+\eta}\sum_{k\in B_l}\left(\frac{w_{kl}}{\xi_{kl}}\right)^{1+\eta}\right]^{\frac{1+\theta}{1+\eta}}}$$

Finally, define the following wage indices:

$$W_m \equiv \left[\sum_z\left(\frac{w_{zm}}{\xi_{zm}}\right)^{1+\eta}\right]^{\frac{1}{1+\eta}}, \quad W \equiv \left[\sum_m\left(\frac{W_m}{\xi_m}\right)^{1+\theta}\right]^{\frac{1}{1+\theta}}$$

Then

$$P^j_{zm} = \frac{\left(\frac{w_{zm}}{\xi_{zm}}\right)^{1+\eta}}{W_m^{1+\eta}} \times \frac{\left(\frac{W_m}{\xi_m}\right)^{1+\theta}}{W^{1+\theta}} = \left(\frac{w_{zm}/\xi_{zm}}{W_m}\right)^{1+\eta} \times \left(\frac{W_m/\xi_m}{W}\right)^{1+\theta} \tag{B.1}$$

With equation B.1 at hand, total labor supplied to firm z in market m can be found by integrating probabilities P^j_{zm} (times $h^j_{zm} = y^j/w_{zm}$ supplied by each worker) over the continuum of workers:

$$l_{zm} = \int_0^1 P^j_{zm}\left(\frac{y^j}{w_{zm}}\right) dF(y) = w_{zm}^{-1} P_{zm} Y \tag{B.2}$$

where $Y \equiv \int_0^1 y^j dF(y)$ is the country-level labor income. To obtain an expression for l_{zm} that is a function of w_{zm}, parameters, and market-level aggregates, I define the following employment indices:

$$L_m \equiv \left[\sum_z (\xi_{zm} l_{km})^{\frac{1+\eta}{\eta}}\right]^{\frac{\eta}{1+\eta}}, \quad L \equiv \left[\sum_m (\xi_m L_m)^{\frac{1+\theta}{\theta}}\right]^{\frac{\theta}{1+\theta}}$$

which, together with equation B.2 and the previously defined wage indices, implies $Y = \sum_{zm} w_{zm} l_{zm} =$

WL and

$$
\begin{aligned}
l_{zm} =& w_{zm}^{-1} P_{zm} Y \\
=& w_{zm}^{-1} \left[\left(\frac{w_{zm}}{\xi_{zm} W_m} \right)^{1+\eta} \times \left(\frac{W_m}{\xi_m W} \right)^{1+\theta} \right] Y \\
=& \cancel{w_{zm}^{-1}} \left[\left(\frac{\cancel{w_{zm}}}{\xi_{zm} \cancel{W_m}} \right) \left(\frac{w_{zm}}{\xi_{zm} W_m} \right)^{\eta} \times \left(\frac{W_m}{\xi_m W} \right)^{\theta} \left(\frac{\cancel{W_m}}{\xi_m \cancel{W}} \right) \right] \cancel{W} L \\
=& (\xi_{zm} \xi_m)^{-1} \left(\frac{w_{zm}/\xi_{zm}}{W_m} \right)^{\eta} \left(\frac{W_m/\xi_m}{W} \right)^{\theta} L
\end{aligned}
$$

Rearranging:

$$
l_{zm} = L \left(\frac{w_{zm}}{W_m} \right)^{\eta} \left(\frac{W_m}{W} \right)^{\theta} \left(\xi_{zm}^{1+\eta} \xi_m^{1+\theta} \right)^{-1} \tag{B.3}
$$

B.2 Other proofs and derivations

B.2.1 Equation 3: $w_{zm} = W \left(\frac{l_{zm}}{L_m} \right)^{\frac{1}{\eta}} \left(\frac{L_m}{L} \right)^{\frac{1}{\theta}} \xi_{zm}^{1+\frac{1}{\eta}} \xi_m^{1+\frac{1}{\theta}}$

The inverse function of the residual labor supply equation 2 (same as Appendix equation B.3) is the wage w_{zm} at which l_{zm} units of labor are supplied to firm z at market m. To check that equation 3 satisfies this criterion, plug it into equation B.2 to obtain the identity $l_{zm} = l_{zm}$. I show this in two steps.

First, plug in the expression for w_{zm} into equation B.2 to get:

$$
\begin{aligned}
l_{zm} =& w_{zm}^{-1} P_{zm} Y \\
=& w_{zm}^{-1} \left(\frac{w_{zm}/\xi_{zm}}{W_m} \right)^{1+\eta} \times \left(\frac{W_m/\xi_m}{W} \right)^{1+\theta} Y \\
=& w_{zm}^{\eta} \left(\frac{1}{W_m \xi_{zm}} \right)^{1+\eta} \times \left(\frac{W_m/\xi_m}{W} \right)^{1+\theta} Y \\
=& \left[W \left(\frac{l_{zm}\xi_{zm}}{L_m} \right)^{\frac{1}{\eta}} \left(\frac{L_m \xi_m}{L} \right)^{\frac{1}{\theta}} \xi_m \xi_{zm} \right]^{\eta} \left(\frac{1/\xi_{zm}}{W_m} \right)^{1+\eta} \times \left(\frac{W_m/\xi_m}{W} \right)^{1+\theta} Y \\
=& \left[W \left(\frac{l_{zm}\xi_{zm}}{L_m} \right)^{\frac{1}{\eta}} \left(\frac{L_m \xi_m}{L} \right)^{\frac{1}{\theta}} \xi_m \xi_{zm} \right]^{\eta} \left(\frac{1/\xi_{zm}}{W_m} \right)^{\eta} \times \left(\frac{W_m/\xi_m}{W} \right)^{\theta} \cancel{W} L \left(\frac{1/\xi_{zm}}{\cancel{W_m}} \right) \times \left(\frac{\cancel{W_m}/\xi_m}{\cancel{W}} \right) \\
=& l_{zm} \xi_m^{\eta} \cancel{\xi_{zm}^{\eta}} \left(\frac{\cancel{\xi_{zm}}}{L_m} \right) \left[W \left(\frac{L_m \xi_m}{L} \right)^{\frac{1}{\theta}} \right]^{\eta} \left(\frac{1/\cancel{\xi_{zm}}}{W_m} \right)^{\eta} \times \left(\frac{W_m/\xi_m}{W} \right)^{\theta} \left(\frac{L}{\cancel{\xi_{zm}}\xi_m} \right) \\
=& l_{zm} \left(\frac{\xi_m^{\eta}}{L_m} \right) \left[W \left(\frac{L_m \xi_m}{L} \right)^{\frac{1}{\theta}} \right]^{\eta} \left(\frac{1}{W_m} \right)^{\eta} \times \left(\frac{W_m/\xi_m}{W} \right)^{\theta} \left(\frac{L}{\xi_m} \right) \\
=& l_{zm} \left(\frac{1}{L_m} \right) \left(\frac{L_m \xi_m}{L} \right)^{\frac{\eta}{\theta}} \left(\frac{W \xi_m}{W_m} \right)^{\eta} \times \left(\frac{W_m/\xi_m}{W} \right)^{\theta} \left(\frac{L}{\xi_m} \right) \\
=& l_{zm} \left(\frac{W \xi_m}{W_m} \right)^{\eta} \times \left(\frac{W_m/\xi_m}{W} \right)^{\theta} \left(\frac{L}{L_m \xi_m} \right) \left(\frac{L_m \xi_m}{L} \right)^{\frac{\eta}{\theta}} \\
=& l_{zm} \left(\frac{W}{W_m/\xi_m} \right)^{\eta} \times \left(\frac{W_m/\xi_m}{W} \right)^{\theta} \left(\frac{L}{L_m \xi_m} \right) \left(\frac{L_m \xi_m}{L} \right)^{\frac{\eta}{\theta}} \\
=& l_{zm} \left(\frac{W_m/\xi_m}{W} \right)^{\theta-\eta} \left(\frac{L_m \xi_m}{L} \right)^{\frac{\eta}{\theta}-1} \\
=& l_{zm} \left(\frac{W}{W_m/\xi_m} \right)^{\eta-\theta} \left(\frac{L_m \xi_m}{L} \right)^{\frac{\eta-\theta}{\theta}} \\
=& l_{zm} \left(\frac{W}{W_m/\xi_m} \right)^{\eta-\theta} \left(\frac{L_m \xi_m}{L} \right)^{\frac{\eta-\theta}{\theta}}
\end{aligned}
$$

Second, I show that $\left(\frac{W}{W_m/\xi_m} \right)^{\eta-\theta} \left(\frac{L_m \xi_m}{L} \right)^{\frac{\eta-\theta}{\theta}} = 1$ by expressing the CES wage index W_m as a function of labor and taste shifters, which can be done by first plugging in the expression for

w_{zm} into the definition of W_m:

$$
\begin{aligned}
W_m &= \left[\sum_{k\in B_n}\left(\frac{w_{zm}}{\xi_{km}}\right)^{1+\eta}\right]^{\frac{1}{1+\eta}} \\
&= \left[\sum_{k\in B_n}\left(\frac{\left[W\left(\frac{l_{km}\xi_{km}}{L_m}\right)^{\frac{1}{\eta}}\left(\frac{L_m\xi_m}{L}\right)^{\frac{1}{\theta}}\xi_m\xi_{km}\right]}{\xi_{km}}\right)^{1+\eta}\right]^{\frac{1}{1+\eta}} \\
&= \left[W^{1+\eta}\left(\frac{1}{L_m}\right)^{\frac{1+\eta}{\eta}}\left[\left(\frac{L_m\xi_m}{L}\right)^{\frac{1}{\theta}}\xi_m\right]^{1+\eta}\sum_{k\in B_n}\left(\frac{\left[(l_{km}\xi_{km})^{\frac{1}{\eta}}\cancel{\xi_{km}}\right]}{\cancel{\xi_{km}}}\right)^{1+\eta}\right]^{\frac{1}{1+\eta}} \\
&= \frac{W}{L_m^{\frac{1}{\eta}}}\left(\frac{L_m\xi_m}{L}\right)^{\frac{1}{\theta}}\xi_m\left[\sum_{k\in B_n}(l_{km}\xi_{km})^{\frac{1+\eta}{\eta}}\right]^{\frac{1}{1+\eta}} \\
&= \frac{W}{\cancel{L_m^{\frac{1}{\eta}}}}\left(\frac{L_m\xi_m}{L}\right)^{\frac{1}{\theta}}\xi_m\cancel{L_m^{\frac{1}{\eta}}} \\
&= W\left(\frac{L_m\xi_m}{L}\right)^{\frac{1}{\theta}}\xi_m
\end{aligned}
$$

Thus, $W_m = W\left(\frac{L_m\xi_m}{L}\right)^{\frac{1}{\theta}}\xi_m$. Recall from the first step that completing the proof requires showing that $\left(\frac{W}{W_m/\xi_m}\right)^{\eta-\theta}\left(\frac{L_m\xi_m}{L}\right)^{\frac{\eta-\theta}{\theta}} = 1$. Plugging in the expression for W_m into this equation gives:

$$
\begin{aligned}
\left(\frac{W}{W_m/\xi_m}\right)^{\eta-\theta}\left(\frac{L_m\xi_m}{L}\right)^{\frac{\eta-\theta}{\theta}} &= \left(\frac{\cancel{W}}{\left[\cancel{W}\left(\frac{L_m\xi_m}{L}\right)^{\frac{1}{\theta}}\cancel{\xi_m}\right]/\cancel{\xi_m}}\right)^{\eta-\theta}\left(\frac{L_m\xi_m}{L}\right)^{\frac{\eta-\theta}{\theta}} \\
&= \left(\frac{L_m\xi_m}{L}\right)^{-\frac{(\eta-\theta)}{\theta}}\left(\frac{L_m\xi_m}{L}\right)^{\frac{\eta-\theta}{\theta}} \\
&= 1
\end{aligned}
$$

which completes the proof that w_{zm} is the inverse function of l_{zm}.

B.2.2 Equation 7: $s_{zm} \equiv \frac{w_{zm}l_{zm}}{\sum_k (w_{km}l_{km})} = \frac{\partial \ln L_m}{\partial \ln l_{zm}}$.

To see why this holds, start from the definition of the labor market index L_m in Section 2 to derive $\partial \ln L_m / \partial \ln l_{zm}$ as

$$\frac{\partial \ln L_m}{\partial \ln l_{zm}} = \frac{(\xi_{km} l_{km})^{\frac{1+\eta}{\eta}}}{\sum_{j=1}^{N_m} (\xi_{jm} l_{jm})^{\frac{1+\eta}{\eta}}}$$

Now set this aside. Plug in equation 3 to the definition $s_{zm} \equiv w_{zm}l_{zm} / \sum_k (w_{km}l_{km})$ to obtain

$$s_{zm} = \frac{(\xi_{km} l_{km})^{\frac{1+\eta}{\eta}}}{\sum_{j=1}^{N_m} (\xi_{jm} l_{jm})^{\frac{1+\eta}{\eta}}}$$

Therefore, $s_{zm} = \partial \ln L_m / \partial \ln l_{zm}$.

B.2.3 Proposition 1: $\mu_m \equiv \frac{\bar{r}_m}{\bar{w}_m} = 1 + \varepsilon_m^{-1} = 1 + \frac{1}{\theta} HHI_m + \frac{1}{\eta}(1 - HHI_m)$,

In this expression, $\bar{w}_m$ and $\bar{r}_m$ are the (employment-weighted) average wage and average marginal revenue product of labor in market m, respectively.

First, I show that $1 + \varepsilon_m^{-1} = 1 + \frac{1}{\theta} HHI_m + \frac{1}{\eta}(1 - HHI_m)$. To see why this holds, let Θ_m denote the set of firms operating in labor market m, and take the (payroll-share-weighted) average of equation 8:

$$\begin{aligned}
\underbrace{\sum_{z \in \Theta_m} s_{zm} \left(1 + \varepsilon_{zm}^{-1}\right)}_{\equiv 1 + \varepsilon_m^{-1}} &= 1 + \sum_{z \in \Theta_m} s_{zm} \left[\frac{1}{\eta}(1 - s_{zm}) + \frac{1}{\theta} s_{zm}\right] \\
&= 1 + \sum_z \left[\frac{1}{\theta} s_{zm}^2 + \frac{1}{\eta}\left(s_{zm} - s_{zm}^2\right)\right] \\
&= 1 + \frac{1}{\theta} HHI_m + \frac{1}{\eta}(1 - HHI_m)
\end{aligned}$$

Second, I show that $\sum_{z \in \Theta_m} s_{zm} \left(1 + \varepsilon_{zm}^{-1}\right) = \frac{\bar{r}_m}{\bar{w}_m}$. To see that this equality holds, aggregate the

firm-level markdown equation $\frac{r_{zm}}{w_{zm}} = 1 + \varepsilon_{zm}^{-1}$ using payroll shares as weights to get:

$$\begin{aligned}
\underbrace{\sum_{z\in\Theta_m} s_{zm}\left(1+\varepsilon_{zm}^{-1}\right)}_{\equiv 1+\varepsilon_m^{-1}} &= \sum_{z\in\Theta_m} s_{zm}\left(\frac{r_{zm}}{w_{zm}}\right) \\
&= \sum_{z\in\Theta_m} \frac{\cancel{w_{zm}} l_{zm}}{\sum_j w_{jm} l_{jm}}\left(\frac{r_{zm}}{\cancel{w_{zm}}}\right) \\
&= \frac{\sum_{z\in\Theta_m} r_{zm} l_{zm}}{\sum_{j\in\Theta_m} w_{jm} l_{jm}} \\
&= \frac{\left(\sum_{z\in\Theta_m} r_{zm} l_{zm}\right) / \left(\sum_{z\in\Theta_m} l_{zm}\right)}{\left(\sum_{j\in\Theta_m} w_{jm} l_{jm}\right) / \left(\sum_{z\in\Theta_m} l_{zm}\right)} \\
&= \frac{\bar{r}_m}{\bar{w}_m} \equiv \mu_m
\end{aligned}$$

B.2.4 Country-level average wage markdown

I show that a particular country-level average of the market-level average wage markdown (i.e., equation 9) equals the country-level average (employment-weighted) wage markdown. The reader can then directly verify that the resulting expression is the inverse of Berger, Herkenhoff and Mongey (2022)'s expression for the "labor market power adjustment" component of the country-level labor share (see authors' equation 10).

Corollary. *Consider the market-level average wage markdown expression from Proposition 1. Then the country-level (employment-weighted) average wage markdown is given by:*

$$\mu \equiv \frac{\bar{r}}{\bar{w}} = \sum_m s_m \mu_m = 1 + \frac{1}{\theta} H\tilde{H}I + \frac{1}{\eta}\left(1 - H\tilde{H}I\right)$$

where $s_m = \frac{\bar{w}_m l_m}{\sum_m \bar{w}_m l_m}$ is market m's payroll share, $H\tilde{H}I = \sum_m s_m HHI_m$ is the country-level payroll-share-weighted average payroll Herfindahl, and $\bar{w}$ and $\bar{r}$ are the (employment-weighted) average wage and average marginal revenue product of labor at the country level, respectively.

Proof. Having provided a more detailed proof for Proposition 1, I use the same steps to show the

country-level aggregation result more directly. In particular:

$$\begin{aligned}
\mu \equiv \frac{\bar{r}}{\bar{w}} =& \frac{\left(\sum_m \bar{r}_m l_m\right) / \left(\sum_m l_m\right)}{\left(\sum_m \bar{w}_m l_m\right) / \left(\sum_m l_m\right)} \\
=& \frac{\sum_m \bar{r}_m l_m}{\sum_m \bar{w}_m l_m} \\
=& \frac{\sum_m \left(\frac{\bar{r}_m}{\bar{w}_m}\right) \bar{w}_m l_m}{\sum_m \bar{w}_m l_m} \\
=& \sum_m \left(\frac{\bar{r}_m}{\bar{w}_m}\right) \frac{\bar{w}_m l_m}{\sum_m \bar{w}_m l_m} \\
=& \sum_m \mu_m s_m \\
=& \sum_m s_m \left[1 + \frac{1}{\theta} HHI_m + \frac{1}{\eta}\left(1 - HHI_m\right)\right] \\
=& 1 + \frac{1}{\theta} \tilde{HHI} + \frac{1}{\eta}\left(1 - \tilde{HHI}\right)
\end{aligned}$$

□

B.2.5 Corollary 1: $\gamma_t \equiv \frac{d\mu_{mt}}{dX} = \left(\frac{1}{\theta} - \frac{1}{\eta}\right)\beta_t$

In this equation, β_t is the effect of an exogenous shock on the payroll Herfindahl. To derive the expression, plug in $\mu_{mt} \equiv 1 + \varepsilon_{mt}^{-1}$ and differentiate:

$$\begin{aligned}
\gamma_t \equiv \frac{d\mu_{mt}}{dX} =& \frac{d\left(1 + \varepsilon_{mt}^{-1}\right)}{dX} \\
=& \left[\frac{d\left(1 + \varepsilon_{mt}^{-1}\right)}{dHHI_{mt}} \cdot \frac{dHHI_{mt}}{dX}\right] \\
=& \left[\frac{d\left(1 + \varepsilon_{mt}^{-1}\right)}{dHHI_{mt}} \cdot \beta_t\right] \\
=& \left(\frac{1}{\theta} - \frac{1}{\eta}\right)\beta_t
\end{aligned}$$

I then compute standard errors for γ_t under the assumption that the effect on concentration and the labor supply parameters are independent. It follows that:

$$\begin{aligned}
\text{Var}(\gamma_t) =& \text{Var}\left[\left(\frac{1}{\theta}-\frac{1}{\eta}\right)\cdot\beta_t\right] \\
=& E\left[\left(\frac{1}{\theta}-\frac{1}{\eta}\right)^2\right]E\left[\beta_t^2\right]-\left[E\left(\frac{1}{\theta}-\frac{1}{\eta}\right)\right]^2\left[E\left(\beta_t\right)\right]^2 \\
=& \left[\text{Var}\left(\frac{1}{\theta}-\frac{1}{\eta}\right)+\left[E\left(\frac{1}{\theta}-\frac{1}{\eta}\right)\right]^2\right]\left[\text{Var}\left(\beta_t\right)+\left[E\left(\beta_t\right)\right]^2\right]-\left[E\left(\frac{1}{\theta}-\frac{1}{\eta}\right)\right]^2\left[E\left(\beta_t\right)\right]^2
\end{aligned}$$

whose components can all be plugged in using sample estimates.

B.2.6 Equation 9 under the setup in BHM

I show that equation 9 holds under the additional assumptions on production function and goods market structure in Berger, Herkenhoff and Mongey (2022), henceforth BHM. In that environment, μ_m should be interpreted as the ratio of the average marginal revenue (net of expenditures in non-labor inputs) to the average wage. I show this in two steps.

To start, consider the environment in BHM. Goods markets are perfectly competitive, with $p_{zm} = 1$ for all firms and markets. Firms compete for labor à la Cournot, solving:

$$\max_{k_{zm},l_{zm}} \pi_{zm} = \underbrace{A_{zm}\left(k_{zm}^{1-\gamma}l_{zm}^{\gamma}\right)^{\alpha}}_{\equiv y_{zm}} - Rk_{zm} - w_{zm}\left(\{l_{zm}, l_{-zm}\}\right)l_{zm} \tag{B.4}$$

where y_{zm} is firm revenues, k_{zm} is capital, A_{zm} is a general firm-market-specific productivity term, R is the rental rate of capital (in perfectly competitive capital markets), and w_{zm} is the wage firm z would have to pay to obtain l_{zm} units of labor, given nested CES labor supply preferences that yield the same expression for ε_{zm}^{-1}, the firm-specific inverse elasticity of residual supply, as derived in Section 2.1.

First, I show that equation 9 holds when the firm optimizes labor *holding capital constant*, denoting this corresponding average wage markdown by $\mu_m^{k-fixed}$. To avoid confusion due to differences in notation, let $mrpl_{zm}^{k-fixed}$ denote BHM's expression for the marginal revenue product of labor of firm z in market m *holding capital constant*, and continue to use the Greek letter μ to denote the wage markdown.[59] Computing the definition of market m's average wage markdown

[59] In BHM, the Greek letter μ refers to the wage take-home share (i.e., the inverse of the wage markdown) holding capital constant.

holding capital constant gives:

$$\mu_m^{k-fixed} \equiv \frac{mrpl_m^{\bar{k}-fixed}}{\bar{w}_m} = \frac{\left(\sum_z mrpl_{zm}^{k-fixed} l_{zm}\right) / \sum_z l_{zm}}{\left(\sum_z w_{zm} l_{zm}\right) / \sum_z l_{zm}} \tag{B.5}$$

$$= \frac{\sum_z \alpha\gamma \left(y_{zm}/l_{zm}\right) l_{zm}}{\sum_z w_{zm} l_{zm}} \tag{B.6}$$

$$= \alpha\gamma \frac{\sum_z y_{zm}}{\sum_z w_{zm} l_{zm}} \tag{B.7}$$

where $mrpl_{zm}^{k-fixed} = \partial y_{zm}/\partial l_{zm}|_k = \alpha\gamma \left(y_{zm}/l_{zm}\right)$.

Simplification of equation B.7 can now be done using the equalities in BHM's Proposition 1.1 (with special care given to note the difference in notation across the two papers). Equation B.7 becomes:

$$\begin{aligned}
\mu_m^{k-fixed} = \alpha\gamma \frac{\sum_z y_{zm}}{\sum_z w_{zm} l_{zm}} &= \alpha\gamma \left[\frac{1}{\alpha\gamma} \sum_z s_{zm} \mu_{zm}^{k-fixed}\right] \\
&= \sum_z s_{zm} \mu_{zm}^{k-fixed} \\
&= \sum_z s_{zm} \left[1 + \frac{1}{\theta} s_{zm} + \frac{1}{\eta}\left(1 - s_{zm}\right)\right] \\
&= 1 + \frac{1}{\theta} HHI_m + \frac{1}{\eta}\left(1 - HHI_m\right)
\end{aligned}$$

where HHI_m is similarly defined as the payroll Herfindahl of labor market m.

Second, I show that, in equilibrium, $\mu_{zm}^{k-fixed} = \mu_{zm}^{k-adjust}$ for all firms z and markets m. In other words, equation 9 holds whether or not optimization of capital is taken into account, *so long as expenditures on capital are netted out of firm revenues*. In this case, $\mu_{zm} = \mu_{zm}^{k-adjust}$ should be interpreted as the ratio of the marginal revenue (net of expenditures on capital) product of labor to the wage.

To show this, note first that since $\mu_{zm} = mrpl_{zm}/w_{zm}$, it suffices to show that in equilibrium $mrpl_{zm}^{k-fixed} = mrpl_{zm}^{k-adjust}$. Letting y_{zm}^{net} denote total firm revenues net of capital expenditures,

it follows that:

$$\begin{aligned} mrpl_{zm}^{k-adjust} \equiv \frac{dy_{zm}^{net}}{dl_{zm}} =& \frac{d\left[f\left(k_{zm}, l_{zm}\right) - Rk_{zm}\right]}{dl_{zm}} \\ =& \frac{\partial f\left(k_{zm}, l_{zm}\right)}{\partial k_{zm}} \frac{dk_{zm}}{dl_{zm}} + \frac{\partial f\left(k_{zm}, l_{zm}\right)}{\partial l_{zm}} - R\frac{dk_{zm}}{dl_{zm}} \\ =& \frac{\partial f\left(k_{zm}, l_{zm}\right)}{\partial l_{zm}} + \frac{dk_{zm}}{dl_{zm}} \underbrace{\left(\frac{\partial f\left(k_{zm}, l_{zm}\right)}{\partial k_{zm}} - R\right)}_{\text{=0 by firm's FOC for } k_{zm}} \\ =& mrpl_{zm}^{k-fixed} \end{aligned}$$

This result, $mrpl_{zm}^{k-fixed} = mrpl_{zm}^{k-adjust}$, also follows directly from the envelope theorem, as the firm is optimizing its non-labor inputs.

B.3 Wage effect decomposition: identities and additional estimates

This appendix states the identities behind the decomposition of Section 7 and reports the estimates that the main text summarizes.

The identity holds for a market's average wage, $\bar{w}_{mt} = \mu_{mt}^{-1}\bar{r}_{mt}$ (equation 9), where μ_{mt}^{-1} is the market average wage take-home share, which depends on the market's payroll Herfindahl, and $\bar{r}_{mt}$ is the market average marginal revenue product of labor. I can therefore decompose the effect of trade liberalization on average wages as:

$$\frac{d\bar{w}_{mt}}{dICE_m} = \frac{d\mu_{mt}^{-1}}{dICE_m}\bar{r}_{mt} + \frac{d\bar{r}_{mt}}{dICE_m}\mu_{mt}^{-1} = -\underbrace{\frac{d\mu_{mt}}{dICE_m}}_{\gamma_t}\frac{1}{\mu_{mt}^2}\bar{r}_{mt} + \frac{d\bar{r}_{mt}}{dICE_m}\mu_{mt}^{-1} \tag{B.8}$$

where $\gamma_t = \left(\frac{1}{\theta} - \frac{1}{\eta}\right)\beta_t$ from Corollary 1 is the effect of changes in import competition exposure on the average wage markdown in local labor market m. The first term is the markdown channel, which Table 4 reports; the second nets changes in the marginal revenue product within firms against changes in the composition of employment across firms, including the firms that enter and exit.

The wage effect that I identify in Section 6 is defined not on the market average, which changes with the firms that enter and exit, but rather on firms present in both years: the market wage premium index ψ_m, the outcome of the reduced form that identifies $1/\theta$, is the 1991-employment-share-weighted mean over incumbent firms of the change in the firm's log wage premium net of its move along its own labor supply curve, $\psi_m = \sum_{z \in S_m} s_{zm0}^e \left[\Delta \ln w_{zm} - \frac{1}{\eta}\Delta \ln l_{zm}\right]$, with S_m the set of incumbents and the shares normalized to sum to one over S_m. For these firms, the wage-setting

equation gives the decomposition firm by firm, and averaging it at baseline shares gives, in logs:

$$\sum_{z\in S_m} s^e_{zm0}\,\Delta\ln w_{zm} = \underbrace{\sum_{z\in S_m} s^e_{zm0}\,\Delta\ln \mu^{-1}_{zm}}_{\text{Via markdowns}} + \underbrace{\sum_{z\in S_m} s^e_{zm0}\,\Delta\ln r_{zm}}_{\text{Via MRPL}} = \psi_m + \frac{1}{\eta}\sum_{z\in S_m} s^e_{zm0}\,\Delta\ln l_{zm}, \qquad \text{(B.9)}$$

where each firm's take-home share μ^{-1}_{zm} follows from its payroll share and the estimated elasticities (equation 8), and its marginal revenue product is its wage premium divided by its take-home share. The identity is exact, so the effect of trade liberalization on incumbents' average wage premium is the sum of its effects on the two terms, and the index differs from that average by incumbents' moves along their labor supply curves.

Across all of a market's firms, the change in the average marginal revenue product, $\bar{r}_{mt} = \sum_z s^e_{zmt} r_{zmt}$, splits into a within-firm and a composition term:

$$\bar{r}_{mt} - \bar{r}_{m0} = \underbrace{\sum_{z\in S_m} s^e_{zm0}\,(r_{zmt} - r_{zm0})}_{\text{Within firms}} + \underbrace{\sum_{z} r_{zmt}\left(s^e_{zmt} - s^e_{zm0}\right)}_{\text{Composition of employment, including entry and exit}}, \qquad \text{(B.10)}$$

where entrants have $s^e_{zm0} = 0$ and exiters $s^e_{zmt} = 0$, and the composition term splits into the change in incumbents' employment shares and the contributions of entrants and exiters. I divide both terms by the logarithmic mean of $\bar{r}_{mt}$ and $\bar{r}_{m0}$, so that they add up to $\Delta\ln\bar{r}_{mt}$ in every market.

The markdown channel and the US benchmark. The market's average take-home share depends on the payroll Herfindahl of all of its firms, including those that exited, and thus falls by more (0.7%) than the average take-home share of incumbent firms (0.38%; Table 5). With the elasticities that BHM estimate for the US, the take-home share would have fallen by 2.8 cents, calculated as $2.8\% \approx \left[\mu_{0,US} + (1/0.42 - 1/10.85) * 0.021\right]^{-1} - \mu^{-1}_{0,US}$, where $\mu_{0,US}$ is the implied pre-liberalization average wage markdown evaluated at Brazil's payroll-weighted average labor market concentration of 0.08 but US elasticities of substitution from BHM ($\theta = 0.42, \eta = 10.85$): $\mu_{0,US} = 1 + (1/0.42) * 0.08 + (1 - 0.08) * (1/10.85)$.

The wage decline at firms facing larger tariff cuts. Within markets, firms whose tariffs fell by more reduced their wage premia relative to their competitors: a one log point larger tariff cut reduced a firm's premium by 0.61 log points (Table 2) and its employment by 0.74. In Appendix Table A.12, I decompose this wage response using the firm-level identity $\Delta\ln w_{zm} = \Delta\ln\mu^{-1}_{zm} + \Delta\ln r_{zm}$, where the firm's take-home share follows from its payroll share and the estimated elasticities (equation 8). Under the weights of Table 2, the market's 1991 employment, the largest markets dominate and a firm's payroll share is close to zero, so these firms' take-home share did not change, and their marginal revenue product of labor fell by the whole 0.61 log points (Panel A). Panel B

reports, for comparison, the effect of the market's exposure on its average firm wage premium and employment, from Table 1. The wage decline at firms facing larger tariff cuts is thus driven by a lower marginal revenue product of labor rather than by greater labor market power. Since firm productivity rose during Brazil's liberalization (Muendler, 2004), and the within-market elasticity is nearly unchanged when I measure tariff shocks as effective rates of protection, which net out input tariff reductions (Appendix E), these MRPL reductions more likely reflect lower output prices than productivity losses, consistent with the pro-competitive effects of import competition predicted by Melitz and Ottaviano (2008) and with the productivity gains documented in India's comparable episode by Topalova and Khandelwal (2011).

Appendix Figure A.11 plots the effect on incumbents' wage premia year by year, with markets weighted by their 1991 formal employment as in Table 1, rather than counted once as in Table 5. After 1991, incumbents' wage premia decline through 1994 and remain lower through 2000 (−0.22 per unit of ΔICE_m, SE of 0.08). Before liberalization, wages in more exposed markets are not stable: they grew faster between 1988 or 1989 and 1991, but not between 1986 or 1990 and 1991. The decline keeps its sign with state-by-year fixed effects and the market's 1986–1990 wage change (−0.16, SE of 0.07) and with microregion-by-year fixed effects (−0.13, SE of 0.08), but these controls do not remove the pre-liberalization movements. The two components of Table 5 are thus not equally well established. The markdown component, 0.4 percentage points, follows from the increase in concentration, which is stable before liberalization (Figure 5), and from the two elasticities. The total wage decline is less certain: across these specifications it ranges from 1.1% to 2.2% per 10% increase in exposure, so that higher markdowns account for between one sixth and one third of it.

C Import Competition Exposure: Theory and Identification

ΔICE_m weights the tariff reductions faced by a market's tradable firms by the firms' contributions to the market's baseline payroll Herfindahl, which determines the market's average markdown (Proposition 1). These are intuitive weights for studying the effects of trade liberalization on concentration: the tariff reductions that count the most are those faced by the firms that account for most of the market's concentration. A reasonable question is how close these weights are to those of the measure of exposure that the model itself implies. In this Appendix, I derive the model's measure and relate ΔICE_m to it.

The derivation requires assumptions that the rest of the paper does not make. The model of Section 2 specifies labor supply and how firms set wages given their labor demand, but it leaves firms' revenue functions unrestricted. It is therefore silent on how much a tariff reduction shifts a firm's labor demand, and on how much the firm's employment responds to that shift. A fully theory-driven, equilibrium-based measure of a market's exposure to trade liberalization, however, must aggregate the tariff reductions of the market's firms according to their effects on the market's equilibrium, which requires taking a stand on both. The assumptions below do so.

Note, however, that my estimates of the paper's key parameters do not depend on these additional assumptions. Per the model, any shock to labor demand that is unrelated to changes in workers' tastes identifies $1/\eta$ and $1/\theta$: the shock need not be a sufficient statistic for the effect of trade liberalization on a local labor market. Proposition 1 and Corollary 1 hold without them as well. The derivation is instead useful for interpreting the effects of ΔICE_m on local labor markets, because it shows what a model-based measure of exposure looks like and how ΔICE_m differs from it.

In the derivation, markdowns respond to the tariff reductions, as in the model of Section 2. The model's measure, ΔICE_m^{ω}, is an average of the tariff reductions faced by the tradable firms of a market, with weights that depend only on the firms' 1991 payroll shares. ΔICE_m coincides with it under conditions that I state, and approximates it otherwise. I also discuss what the derivation implies for concentration.

I then discuss how ΔICE_m and the model's measure relate to the most commonly used measure of exposure in analyses of Brazil's trade liberalization, the RTR (regional tariff reductions) of Kovak (2013) and Dix-Carneiro and Kovak (2017). Finally, I document which variation in ΔICE_m identifies my market-level estimates, and compare it with the variation in the RTR. Proofs are at the end of this Appendix.

C.1 The model's measure of exposure

This section derives ΔICE_m^{ω}, the model's measure of market m's exposure to the tariff reductions: the average of the tariff reductions faced by the market's tradable firms to which the changes in the

market's labor supply and wage indices are proportional. I proceed in two steps. Proposition 2 shows how a market's indices respond to any labor demand shocks at its firms. It uses Assumptions 1 and 2, which are consistent with the model of Section 2 and narrow it to a case in which this response has a closed form. Corollary 2 then obtains ΔICE_m^ω with Assumptions 3 to 5, which are new: they specify how tariff reductions shift firms' labor demand, which Section 2 leaves unspecified. The next section compares ΔICE_m^ω with ΔICE_m.

I use the model of Section 2 as is: labor supply (equation 3), wage setting (equation 5), and markdowns μ_{zm} (equation 8). Let $r_{zm} \equiv \partial R_z / \partial l_{zm}$ denote firm z's marginal revenue product of labor in market m. I compare the equilibria before and after the tariff reductions, to first order. Wages, employment, payroll shares, and markdowns all change between the two equilibria; the payroll shares s_{zm} that appear in the weights below are those of 1991.

Assumption 1 (No other changes). The taste shifters ξ_{zm} and ξ_m, the country-level indices W and L, and the set of firms operating in each market do not change with tariffs.

Assumption 2 (Labor demand). The marginal revenue product of labor of firm z in market m satisfies

$$d \ln r_{zm} = x_{zm} - \frac{1}{e_{zm}} d \ln l_{zm}, \tag{C.1}$$

where x_{zm} is the firm's labor demand shock, the change in its marginal revenue product at a given level of employment, and $e_{zm} > 0$ is the elasticity of its labor demand.

The first part of Assumption 1 is the exclusion restriction of Section 5. Its second part is a normalization when the elasticity of labor demand is the same for all firms: the country-level terms are then common to all markets and absorbed by the constants of the market-level regressions. Its third part excludes the entry and exit of firms, and with them the changes in the set of markets in which a firm operates that Section 2.2 allows. Assumption 2 restricts the revenue function R_z of equation 4, on which the rest of the paper remains agnostic: the marginal revenue product of labor in market m depends neither on the employment of the other firms in that market, as when the firms that share a local labor market sell their output in national or world markets, nor on firm z's employment in its other markets.

Proposition 2. *Under Assumptions 1 and 2,*

$$d \ln L_m = \frac{\sum_{z \in \Theta_m} \omega_{zm} \, x_{zm}}{\sum_{z \in \Theta_m} \omega_{zm} / e_{zm} + \frac{1}{\theta}}, \qquad d \ln W_m = \frac{1}{\theta} \, d \ln L_m, \tag{C.2}$$

where

$$\omega_{zm} \equiv \frac{s_{zm}/D_{zm}}{\sum_{k \in \Theta_m} s_{km}/D_{km}}, \qquad D_{zm} \equiv \frac{1}{e_{zm}} + \frac{1}{\eta} + \left(\frac{1}{\theta} - \frac{1}{\eta}\right)\left(1 + \frac{1}{\eta}\right) \frac{s_{zm}}{\mu_{zm}}, \tag{C.3}$$

provided that $D_{zm} > 0$ for all firms, which holds when $1/\theta \geq 1/\eta$.

Proof. See Appendix C.6. □

The labor demand shocks of a market's firms affect the market's labor supply and wage indices only through their weighted average. A firm's weight is its payroll share, its marginal effect on the market's labor supply index (equation 7), divided by D_{zm}, whose inverse is the response of the firm's employment to its own shock for a given L_m: smaller when the firm's labor demand is less elastic, when labor supply to the firm is less elastic, and when the firm's markdown rises with its payroll share. Since the markdown μ_{zm} is a function of the payroll share, the weights depend only on the firms' 1991 payroll shares, given the elasticities.

For the second step, let $\Delta\tau_z \equiv \Delta \ln(1 + \tau_{i(z)})$ denote the tariff reduction faced by firm z (equation 17), positive for a tariff cut, and Θ_m^T the tradable firms of market m, as in Section 4.1.

Assumption 3 (Pass-through). For tradable firms, $x_{zm} = -\psi \, \Delta\tau_z$, where $\psi > 0$ is the pass-through of tariff reductions to labor demand, common to all firms.

Assumption 4 (Common elasticity). The elasticity of labor demand is the same for all firms, $e_{zm} = e$.

Assumption 5 (Non-tradable firms). For firms outside Θ_m^T, $x_{zm} = \bar{x}_m^T \equiv \sum_{k \in \Theta_m^T} (\omega_{km}/\omega_m^T) \, x_{km}$, the average labor demand shock of the market's tradable firms with the weights of equation C.3, where $\omega_m^T \equiv \sum_{k \in \Theta_m^T} \omega_{km}$, with $x_{zm} = 0$ in markets without tradable firms.

Assumption 5 specifies the shocks of non-tradable firms, whose output demand is local, without modeling local goods markets.

Corollary 2. *Under Assumptions 1 to 5, $d \ln W_m = \frac{1}{\theta} \, d \ln L_m$ and*

$$d \ln L_m = -\frac{\psi}{\frac{1}{e} + \frac{1}{\theta}} \, \Delta ICE_m^{\omega}, \tag{C.4}$$

where

$$\Delta ICE_m^{\omega} \equiv \sum_{z \in \Theta_m^T} \frac{\omega_{zm}}{\omega_m^T} \, \Delta\tau_z \tag{C.5}$$

in markets with tradable firms, and $\Delta ICE_m^{\omega} \equiv 0$ in markets without them.

Proof. See Appendix C.6. □

By Corollary 2, ΔICE_m^{ω} is the model's measure of the market's exposure: the changes in the market's wage and labor supply indices are proportional to it, with the same factor in all markets. Equation C.4 is the model's counterpart of the first stage in equation 20. With a common elasticity of labor demand, D_{zm} varies across the firms of a market only with s_{zm}/μ_{zm}, so that when $1/\theta > 1/\eta$ a firm's weight increases less than proportionally with its payroll share.

C.2 Relation to ΔICE_m

Let $s_m^T \equiv \sum_{z \in \Theta_m^T} s_{zm}$ denote the share of tradable firms in the market's payroll. The payroll-share-weighted average of the tariff reductions of the market's tradable firms,

$$\Delta ICE_m^{\text{lin}} \equiv \sum_{z \in \Theta_m^T} \frac{s_{zm}}{s_m^T} \Delta\tau_z, \tag{C.6}$$

replaces D_{zm} with a value common to the firms of a market. It equals ΔICE_m^{ω} when $1/\theta = 1/\eta$, when the tradable firms of a market have equal payroll shares, or when they face the same tariff reduction, and it approximates it otherwise: at the estimates of Section 6, the term of D_{zm} that varies with the payroll share is 0.06 for a firm with a payroll share of 10% and 0.27 for a share of 50%, against $1/e + 1/\eta \geq 0.83$. The measure of Section 4, ΔICE_m of equation 11, averages the same tariff reductions over the same firms with weights κ_{zm} proportional to squared payroll shares, and

$$\Delta ICE_m = \Delta ICE_m^{\text{lin}} + \frac{s_m^T}{HHI_m^T} \text{Cov}_m^T(s_{zm}, \Delta\tau_z), \tag{C.7}$$

where $HHI_m^T \equiv \sum_{z \in \Theta_m^T} s_{zm}^2$ and Cov_m^T is the payroll-share-weighted covariance across the tradable firms of market m.

Assumption 6 (Tariff reductions and firm size). Across the tradable firms of each market, tariff reductions are uncorrelated with payroll shares: $\text{Cov}_m^T(s_{zm}, \Delta\tau_z) = 0$.

Under Assumption 6, $\Delta ICE_m = \Delta ICE_m^{\text{lin}}$; in markets whose tradable firms face the same tariff reduction, the three measures coincide. Where Assumption 6 does not hold, ΔICE_m gives the tariff reductions of the market's larger tradable firms more weight than their payroll shares, whereas the model's measure gives them somewhat less. ΔICE_m is thus an approximation to the model's measure of exposure, not the measure itself. Column (2) of Appendix Table A.4 reports the estimates of Section 4 with $\Delta ICE_m^{\text{lin}}$ in place of ΔICE_m.

C.3 Concentration

Proposition 3. *Under Assumptions 1 and 2,*

$$dHHI_m = 2\left(1+\frac{1}{\eta}\right)\sum_{z\in\Theta_m}\frac{s_{zm}^2}{D_{zm}}\left[x_{zm}-\left(\frac{1}{e_{zm}}+\frac{1}{\theta}\right)d\ln L_m\right], \tag{C.8}$$

provided that $D_{zm} > 0$ *for all firms.*

Proof. See Appendix C.6. □

Concentration rises when employment falls by less at firms with larger payroll shares. With a common pass-through and elasticity of labor demand (Assumptions 3 to 5), equation C.8 becomes

$$dHHI_m = -2\psi\left(1+\frac{1}{\eta}\right)\sum_{z\in\Theta_m^T}\frac{s_{zm}^2}{D_{zm}}\left(\Delta\tau_z-\Delta ICE_m^{\omega}\right):$$

a tariff reduction common to the tradable firms of a market leaves the Herfindahl unchanged, and concentration rises when the tradable firms with larger payroll shares face smaller tariff reductions than the market's average. The increase in concentration estimated in Section 4 can also arise from departures from these assumptions. One is firm heterogeneity in pass-through or in labor demand elasticities, which Propositions 2 and 3 accommodate: a lower pass-through at larger firms, for example because they export part of their output, or less elastic labor demand at larger firms, as with the linear demand of Melitz and Ottaviano (2008). The other is the entry and exit of firms, which Assumption 1 excludes and which Section 4.3 shows to be important.

C.4 Relation to Kovak (2013)

I now relate the model's measure to the exposure measure of Kovak (2013), also used by Dix-Carneiro and Kovak (2017): RTR (regional tariff reductions). In Kovak (2013), a region produces goods i with labor and industry-specific factors in fixed supply; labor is mobile across the region's industries, immobile across regions, and supplied inelastically; technology has constant returns to scale, all markets are perfectly competitive, and the region takes the prices P_i of tradable goods as given. Clearing of the regional labor market implies

$$d\ln w = \sum_i \beta_i\, d\ln P_i, \qquad \beta_i \equiv \frac{\lambda_i\, e_i}{\sum_{i'}\lambda_{i'}\, e_{i'}}, \tag{C.9}$$

where w is the regional wage, λ_i is industry i's share of regional employment, and e_i is the elasticity of the industry's labor demand. Kovak (2013) measures price changes of tradable goods

with tariff changes, $d \ln P_i = d \ln(1 + \tau_i)$, drops the non-tradable sector, whose price moves with the region's tradable prices under identical Cobb-Douglas preferences, and rescales the weights of tradable industries to sum to one. The result, RTR, is his model's measure of the region's exposure to liberalization.

ΔICE_m^ω has the same form, with firms in place of industries. In both models, exposure is a weighted average of tariff reductions, and the weight of each unit is the product of two terms: its size in the labor market, and how much its employment responds to a shock to its own labor demand. In equation C.9, these are the industry's share of regional employment, λ_i, and the elasticity of its labor demand, e_i: the tariff reduction of an industry counts more when the industry employs more of the region's workers, and when its employment falls by more for a given decline in its price. In equation C.3, they are the firm's payroll share, s_{zm}, which is the firm's marginal effect on the market's labor supply index (equation 7), and $1/D_{zm}$, the response of the firm's employment to its own labor demand shock.

The two models differ in what determines this response. In Kovak (2013), an industry hires any number of workers at the regional wage, so its employment response depends only on the elasticity of its labor demand. In my model, a firm faces an upward-sloping labor supply curve and sets its wage, which adds two terms to D_{zm}. The first, $1/\eta$, reflects that a firm that reduces employment moves down its labor supply curve: its wage falls, which dampens the decline in its employment. The second, $\left(\frac{1}{\theta} - \frac{1}{\eta}\right)\left(1 + \frac{1}{\eta}\right)\frac{s_{zm}}{\mu_{zm}}$, reflects that the firm's markdown increases with its payroll share when $1/\theta > 1/\eta$ (equation 8): a firm that loses payroll share marks wages down by less, which further dampens the decline in its employment, the more so the larger the firm. This second term is why the weights of ΔICE_m^ω rise less than proportionally with payroll shares.

The two models also differ in how a market absorbs the average shock to its firms. In Kovak (2013), labor does not move across regions, so the regional wage absorbs the shock in full. In my model, workers substitute across local labor markets with a finite elasticity, so the shock moves both the market's labor supply index and its wage index, in the proportion $1/\theta$ (equation C.2).

Equation C.9 is the limit of Proposition 2 in which these differences vanish. When firms take wages as given, the markdown term drops out and $D_{zm} = 1/e_{zm} + 1/\eta$. If, in addition, workers view the firms of a market as identical ($1/\eta \to 0$, $\xi_{zm} = 1$), then $D_{zm} = 1/e_{zm}$ and payroll shares equal employment shares; and if labor does not move across markets ($1/\theta \to \infty$), $d \ln W_m = \sum_z s_{zm} e_{zm} x_{zm} \Big/ \sum_k s_{km} e_{km}$, which is equation C.9 with firms in place of industries, where x_{zm} is the change in the price of the industry's good and e_{zm} is e_i. The assumptions that take me from Proposition 2 to ΔICE_m^ω also have counterparts in Kovak (2013): Assumption 3 corresponds to his measurement of price changes with tariff changes, and Assumption 5, when markets are microregions, to his result for the price of the non-tradable good. Assumption 4 is more restrictive than his treatment of labor demand: he lets the elasticity of labor demand differ across industries and

calibrates it with each industry's factor shares, whereas I assume an elasticity common to all firms.

The two measures therefore differ in what they average: RTR averages tariff reductions over the industries of a region, with weights given by the industries' shares of the region's employment and their labor demand elasticities, and ΔICE_m over the tradable formal firms of a local labor market, with weights given by their payroll shares. Appendix C.5 documents how these differences translate into different sources of identifying variation.

C.5 Sources of identifying variation

ΔICE_m is an alternative to the exposure measure of Kovak (2013), also used by Dix-Carneiro and Kovak (2017): RTR (regional tariff reductions). In this section, I use tools for shift-share designs (Goldsmith-Pinkham, Sorkin and Swift, 2020; Borusyak, Hull and Jaravel, 2022; Borusyak and Hull, 2023) to identify which variation in ΔICE_m drives each of my market-level estimates, and thus the assumption under which each can be interpreted as a causal effect. I also report the effects of the liberalization under the RTR, using the specification of Dix-Carneiro and Kovak (2017), in my RAIS data. All estimates are relative effects: equation (12) compares more with less exposed markets, and the year fixed effects absorb effects common to all markets. Unless otherwise noted, I weight local labor markets by their 1991 formal employment, as in my main specification; Appendix D discusses the unweighted estimates. In short, I find that the effects of ΔICE_m on concentration and firm exit are identified by differences in tariff cuts across sectors, and that its effect on employment rests more on the contrast between markets with and without tradable firms in 1991. The RTR varies across regions mainly with the share of agriculture, and the estimates under the RTR do not stand out from those that I obtain when I reassign tariff cuts at random across industries.

What ΔICE_m compares

ΔICE_m averages the tariff cuts faced by a market's tradable firms, weighted by their squared payroll shares, so it varies with which sectors those firms are in and with how the market's payroll is distributed among them. It is unrelated to the microregion's 1991 share of agriculture and mining in employment. It is zero in the 15 percent of microregion $\times$ occupation markets without a tradable firm in 1991. I call these markets never-exposed and the others exposed.

To separate the sources of variation in ΔICE_m, I use three versions of it: (a) ΔICE_m; (b) ΔICE_m recentered on its expected value when I reassign tariff cuts at random across CNAE sectors within IBGE subsectors (Borusyak and Hull, 2023), which keeps only differences in tariff cuts across detailed sectors; and (c) an indicator for exposed markets, which keeps only the contrast with never-exposed markets. An effect identified by the tariff cuts survives in (b). An effect identified by composition is close to zero in (b) and appears in (c). In Tables 1 and C.2, I implement (b) with the

recentered measure as the regressor. Controlling instead for the expected value of ΔICE_m interacted with year, an alternative implementation of the same correction (Borusyak and Hull, 2023), gives 0.151 against 0.168 for the effect on concentration (Table D.3).

Concentration and firm exit under ΔICE_m

My estimates of the effects on concentration and firm exit are identified by the tariff cuts. Recentering ΔICE_m on its expected value leaves my estimates nearly unchanged: by 2000 the Herfindahl rises by 0.168 against 0.172, and the number of firms falls by 1.09 against 0.97 log points (Table 1, column (2)). When I reassign the 1990–1994 tariff changes at random across CNAE sectors, holding each market's firms and shares fixed, the placebo estimates center near zero, and the actual estimate exceeds all but 1.6 percent of them (Figure C.2, panel (a)). The effect changes little with the RTR or the microregion's agriculture share, interacted with year, as controls (0.130 and 0.126 against 0.172; Table C.1). Interpreting it as causal requires that tariff cuts be as good as randomly assigned across sectors (Borusyak, Hull and Jaravel, 2022). The Herfindahl rises by 0.054 in 1992 and by 0.129 by 1995, three quarters of its 2000 level, whereas formal employment falls only after 1995: concentration increases while tariffs are being cut, and the decline in employment follows.

Table C.1: Controls for agriculture intensity and tradable sector intensity

	Δ Payroll Herfindahl (premium-based) (1)	Δ Employment Herfindahl (2)	Δ Log formal employment (3)
Main specification (Table 1)	0.172	0.109	-0.398
	(0.042)	(0.024)	(0.186)
Control: 1991 census agricultural share × year	0.130	0.071	-0.201
	(0.040)	(0.021)	(0.163)
ΔICE without agricultural firms	0.162	0.096	-0.299
	(0.042)	(0.024)	(0.184)
Dropping top-quartile agricultural microregions	0.163	0.097	-0.411
	(0.043)	(0.024)	(0.190)
Control: Dix-Carneiro and Kovak (2017)'s RTR × year	0.126	0.068	-0.207
	(0.040)	(0.020)	(0.167)
Control: 1991 payroll share of tariff firms × year	0.180	0.116	-0.313
	(0.042)	(0.024)	(0.175)
Control: 1{tariff firms in 1991} × year	0.162	0.093	-0.390
	(0.044)	(0.025)	(0.192)
Markets with tariff firms only	0.162	0.093	-0.390
	(0.044)	(0.025)	(0.192)
Microregion × year fixed effects	0.084	0.044	-0.027
	(0.035)	(0.021)	(0.150)
Occupation × year and size decile × year FE	0.128	0.067	-0.195
	(0.040)	(0.021)	(0.142)

Notes: Standard errors in parentheses, clustered at the local labor market level. All entries report the 2000 coefficient on ΔICE_m from the stacked differences-in-differences regressions of Table 1, weighted by the local labor market's 1991 formal employment. The agricultural employment share is the microregion's 1991 share of tradable employment in agriculture, from the census. The RTR is Dix-Carneiro and Kovak (2017)'s RTR exposure measure, computed from the 1991 census. The tariff firms' payroll share is the 1991 payroll share of the market's firms in sectors with a tariff. ΔICE_m without agricultural firms excludes agricultural firms from the market's tariff firms. The top quartile of agricultural share is defined across microregions.

Employment and wages under ΔICE_m

Employment and wages. Formal employment falls in more exposed markets by 2000, mostly after 1994; the formal wage premium changes little (−0.105, SE of 0.089, per unit of ΔICE_m; Table 1), and its 1986–1989 coefficients show no trend (p-value of 0.31; Table D.3).

Source of variation and identifying assumption. The employment effect rests more on 1991 composition than the effects on concentration and on the number of firms. With the recentered measure it is −0.30 (SE of 0.21), against −0.40 (SE of 0.19) for ΔICE_m (Table 1, column (2)). To the extent that it rests on that contrast, reading it as causal requires that exposed and never-exposed markets would have followed the same paths without the liberalization (Goldsmith-Pinkham, Sorkin and Swift, 2020). Its pre-period coefficients differ from zero (Table D.3), and the removal of import licenses in 1990, the 1990–1992 recession, the 1994 Real Plan and the 1999 devaluation all fell on

the same contrast between tradable and other sectors; this is why the estimation of $1/\theta$ includes state fixed effects and 1986–1990 trends, as in Dix-Carneiro and Kovak (2017). State × year fixed effects and the market's 1986–1990 trend halve the employment effect, to −0.194 (SE of 0.145), while the effects on concentration and on the number of firms remain statistically significant (Appendix Table A.5). The effect on the wage premium changes little when I recenter ΔICE_m (−0.119 against −0.105; Table 1) and is imprecise in both versions.

The wage premium index. Table C.2 applies the same three versions of ΔICE_m to the two objects of the $1/\theta$ estimation, on its own design (1991–2000 long differences, state fixed effects and 1986–1990 trends): the market wage premium index ψ_m, the reduced form of Table 3, and the market employment index, its first stage. Unweighted, the index falls by 0.183 (SE of 0.050) per unit of ΔICE_m, by 0.106 (SE of 0.063) with the recentered measure, and by 0.103 (SE of 0.057) among exposed markets alone; weighted, by 0.214 (SE of 0.065), 0.122 (SE of 0.077) and 0.217 (SE of 0.067). The wage decline at incumbent firms is thus identified in part by the differences in tariff cuts across sectors. The employment index is not: its first stage is −0.294 (SE of 0.044) unweighted and −0.008 (SE of 0.060) with the recentered measure, so it comes from the contrast between markets with and without tradable firms, and $1/\theta$, the ratio of the two, rests on that contrast under the parallel-trends assumption that the pre-period controls address.

Table C.2: The wage premium index and the employment index by source of variation in ΔICE_m

		ΔICE_m	Recentered ΔICE_m	Exposed indicator	ΔICE_m, exposed indicator as control	ΔICE_m, exposed markets only
		(1)	(2)	(3)	(4)	(5)
Δ Market wage premium index, ψ_m (reduced form)	Unweighted	-0.183	-0.106	-0.054	-0.094	-0.103
		(0.050)	(0.063)	(0.014)	(0.057)	(0.057)
	Weighted	-0.214	-0.122	-0.023	-0.217	-0.217
		(0.065)	(0.077)	(0.035)	(0.067)	(0.068)
Δ Market employment index (first stage)	Unweighted	-0.294	-0.008	-0.154	0.076	0.078
		(0.044)	(0.060)	(0.012)	(0.052)	(0.052)
	Weighted	0.013	0.066	-0.130	0.051	0.050
		(0.104)	(0.124)	(0.030)	(0.109)	(0.109)
State FE and 1986–1990 trends		Yes	Yes	Yes	Yes	Yes
Local labor markets		17,628	15,058	17,628	17,628	15,058

Notes: Standard errors in parentheses, clustered by local labor market. Each cell reports the coefficient on the column's regressor in a regression of the 1991–2000 change in the row's object on that regressor, with state fixed effects and the market's 1986–1990 changes in log employment and wage premium, as in Table 3; unweighted, or weighted by the market's 1991 formal employment. The market wage premium index ψ_m is the outcome of the reduced form of Table 3, and the market employment index the outcome of its first stage (Section 6). Column (2) subtracts from ΔICE_m its expected value when the 1990–1994 tariff changes are reassigned at random across CNAE sectors within IBGE subsectors (Borusyak and Hull, 2023), on the markets with a tariff firm in 1991. Column (3) replaces ΔICE_m with an indicator for these markets; column (4) adds the indicator as a control; column (5) restricts the sample to them.

ΔICE_m and the RTR

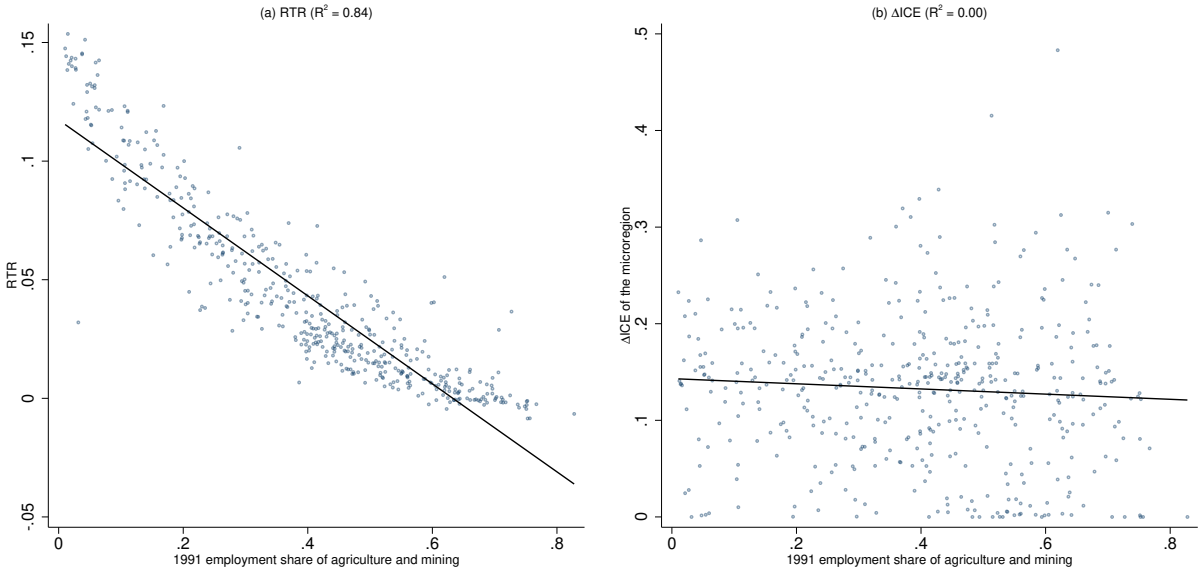
With the specification of Dix-Carneiro and Kovak (2017) at the microregion level, the RTR as the measure of exposure, state fixed effects and 1986–1990 trends, regional formal earnings premia

estimated in RAIS as in their paper give −0.442 (SE of 0.140) by 2000, against their −0.529 (SE of 0.141) (Table C.3; their Table 2). The data of this paper thus reproduce their findings under their design; the two measures differ in what they compare across regions.

The random reassignment of tariff cuts that I apply to ΔICE_m (Figure C.2, panel (a)) gives a different result for the RTR. When I reassign the 1990–1995 tariff changes at random across the tradable industries of Dix-Carneiro and Kovak (2017), holding each microregion's 1991 industry composition fixed, and re-estimate their specification for formal earnings premia, the placebo estimates are widely dispersed around zero, and 88 percent of them are larger in absolute value than the actual estimate of −0.442 (Figure C.2, panel (b)).

What the two designs identify. Dix-Carneiro and Kovak (2017)'s RTR exposure measure and ΔICE_m both rest on a composition contrast for the wage and employment effects as of 2000: for the RTR, the microregion's industry composition, in which the share of agriculture, an industry with a small tariff cut, dominates the cross-regional variation (a regression of the RTR on the census agricultural share of tradable employment has an R^2 of 0.93, and of 0.84 on the employment share of agriculture and mining, Figure C.1; agriculture is 55% of census tradable employment and 6% of RAIS employment); for ΔICE_m, whether the market had tradable firms in 1991. Both designs assess the contrast with 1986–1990 trends. ΔICE_m is unrelated to the agricultural share (an R^2 of 0.0002 in the same regression), so the two measures rank regions differently and their wage effects are not estimates of the same comparison, and under ΔICE_m part of the decline in the wage premium index survives recentering (Table C.2). The two designs answer different questions. The RTR asks how a region whose workers were in industries with larger tariff cuts fared relative to one whose workers were in industries with smaller cuts, agriculture above all; ΔICE_m asks how a formal labor market whose employers faced tariff cuts fared relative to one whose employers faced none. Under their design I reproduce their earnings decline in RAIS; under mine, incumbent firms' wage premia fall, by an amount of the same sign, and I estimate the decline in the market average imprecisely under my main market-level specification (Table 1). The effects on concentration and firm exit, which are the paper's subject, are identified by the tariff cuts themselves and do not depend on this contrast: with tariffs measured for 285 sectors and firms' payroll shares as the shares, ΔICE_m varies across markets with tradable firms only with the tariff cuts their firms faced.

Figure C.1: Dix-Carneiro and Kovak (2017)'s RTR exposure measure and ΔICE_m against the microregion's 1991 agricultural employment share

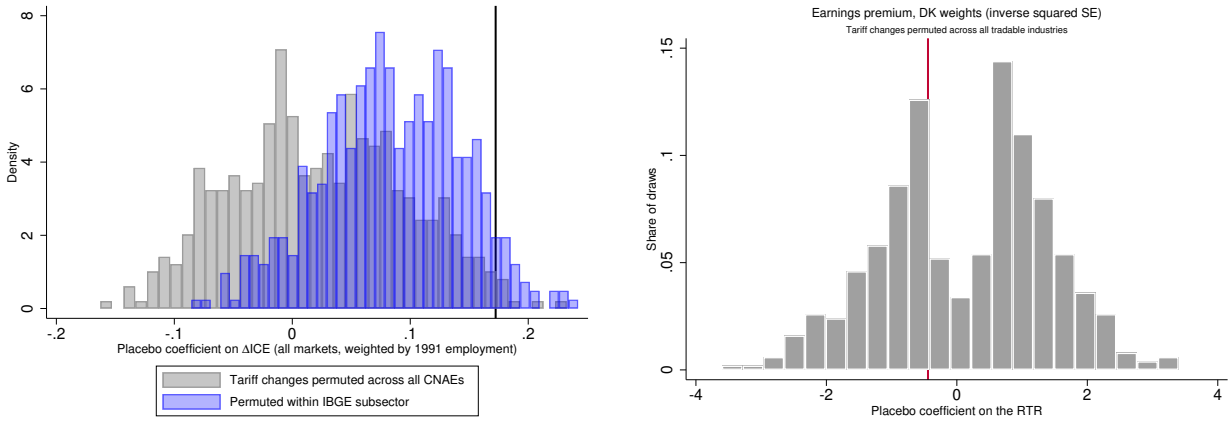


Notes: Each dot is a microregion. The horizontal axis is the 1991 share of agriculture and mining in the microregion's employment, from the census. Panel (a) plots Dix-Carneiro and Kovak (2017)'s RTR exposure measure, computed from the 1991 census (R^2 of the fitted line, 0.84); panel (b) plots ΔICE_m with the microregion as the local labor market (R^2 of 0.00). Lines are OLS fits.

Figure C.2: Placebo estimates under random reassignment of tariff cuts: ΔICE_m and Dix-Carneiro and Kovak (2017)'s RTR exposure measure

(a) ΔICE_m: payroll Herfindahl

(b) RTR: formal earnings premia



Notes: Panel (a) plots the 2000 coefficients of equation (12) for the premium-based payroll Herfindahl, with markets weighted by their 1991 formal employment, when ΔICE_m is recomputed after reassigning the 1990–1994 tariff changes at random across all CNAE sectors (gray) or across CNAE sectors within IBGE subsectors (blue), holding each market's firms and payroll shares fixed. Panel (b) plots the coefficients on Dix-Carneiro and Kovak (2017)'s RTR exposure measure in their specification for the 1991–2000 change in microregions' formal earnings premia (Table C.3, column (3)) when the RTR is recomputed after reassigning the 1990–1995 tariff changes at random across tradable industries, holding each microregion's 1991 industry composition fixed (500 draws). The vertical lines are the estimates with the actual tariff changes.

Table C.3: Replication of Dix-Carneiro and Kovak (2017) in RAIS: effects of their RTR exposure measure on microregions, 1991–2000

Change in outcome, 1991–2000:	(1)	(2)	(3)
Panel A. Log formal earnings premia			
Dix-Carneiro and Kovak (2017)'s RTR exposure measure	-0.353	-0.566	-0.442
	(0.159)	(0.162)	(0.140)
Panel B. Log formal employment			
Dix-Carneiro and Kovak (2017)'s RTR exposure measure	-3.752	-3.570	-3.555
	(0.512)	(0.554)	(0.576)
State fixed effects (26)		Yes	Yes
Pre-trend, 1986–1990			Yes
Microregions	475	474	474

Notes: Microregion-level regressions of the 1991–2000 change in the outcome on Dix-Carneiro and Kovak (2017)'s RTR exposure measure, computed from the 1991 census following their replication code. Panel A uses the microregion's log formal earnings premium, estimated in RAIS for each year as in Dix-Carneiro and Kovak (2017), from a regression of workers' log December earnings on age, sex, and education, industry fixed effects, and microregion fixed effects; regressions in Panel A are weighted by the inverse of the squared standard error of the change in the premium. Panel B uses the microregion's log formal employment, unweighted. Column (2) adds state fixed effects; column (3) also adds the microregion's 1986–1990 change in the outcome, the specification of Dix-Carneiro and Kovak (2017). Standard errors clustered by mesoregion in parentheses.

C.6 Proofs

Propositions 2 and 3 (equation C.8). Log-differentiating the first-order condition $r_{zm} = w_{zm}\mu_{zm}$ gives $d\ln r_{zm} = d\ln w_{zm} + d\ln \mu_{zm}$. By Assumption 1 and equations 3, 7, and 8,

$$d\ln w_{zm} = \frac{1}{\eta}\, d\ln l_{zm} + \left(\frac{1}{\theta} - \frac{1}{\eta}\right) d\ln L_m, \qquad d\ln L_m = \sum_{k\in\Theta_m} s_{km}\, d\ln l_{km},$$

$$d\ln \mu_{zm} = \left(\frac{1}{\theta} - \frac{1}{\eta}\right)\frac{ds_{zm}}{\mu_{zm}}, \qquad ds_{zm} = \left(1 + \frac{1}{\eta}\right) s_{zm}\left(d\ln l_{zm} - d\ln L_m\right),$$

where the last equality uses $s_{zm} = (\xi_{zm} l_{zm})^{\frac{1+\eta}{\eta}} \,/\, \sum_{k\in\Theta_m}(\xi_{km} l_{km})^{\frac{1+\eta}{\eta}}$ (Appendix B.2). Substituting these expressions and Assumption 2 into the differentiated first-order condition and rearranging gives

$$d\ln l_{zm} - d\ln L_m = \frac{1}{D_{zm}}\left[x_{zm} - \left(\frac{1}{e_{zm}} + \frac{1}{\theta}\right) d\ln L_m\right]. \tag{C.10}$$

Multiplying equation C.10 by s_{zm} and summing over the firms of the market gives zero on the left-hand side, and solving for $d\ln L_m$ gives the first part of equation C.2; the second part follows from $W_m = W\,(L_m\xi_m/L)^{\frac{1}{\theta}}\,\xi_m$ (Appendix B.2). For equation C.8, $dHHI_m = 2\sum_z s_{zm}\, ds_{zm} = 2\left(1+\frac{1}{\eta}\right)\sum_z s_{zm}^2\,(d\ln l_{zm} - d\ln L_m)$, and the result follows from equation C.10.

Corollary 2, the Herfindahl under Assumptions 3 to 5, and equation C.7. Under Assumption 4, $\sum_z \omega_{zm}/e_{zm} = 1/e$, so that equation C.2 becomes $(1/e + 1/\theta)\, d\ln L_m = \bar{x}_m \equiv \sum_{z\in\Theta_m}\omega_{zm}x_{zm}$.

By Assumption 5, $\bar{x}_m = \omega_m^T \bar{x}_m^T + (1 - \omega_m^T)\, \bar{x}_m^T = \bar{x}_m^T$, and by Assumption 3, $\bar{x}_m^T = -\psi\, \Delta ICE_m^\omega$, which gives equation C.4. Substituting $(1/e + 1/\theta)\, d \ln L_m = \bar{x}_m^T$ into equation C.8, the term in brackets is $x_{zm} - \bar{x}_m^T$, which is zero for non-tradable firms and $-\psi(\Delta\tau_z - \Delta ICE_m^\omega)$ for tradable firms. For equation C.7, the weights of equation 11 are $\kappa_{zm} = s_{zm}^2 / HHI_m^T$, so that $\frac{s_m^T}{HHI_m^T} \mathrm{Cov}_m^T(s_{zm}, \Delta\tau_z) = \frac{s_m^T}{HHI_m^T} \left[\sum_{z \in \Theta_m^T} \frac{s_{zm}^2}{s_m^T} \Delta\tau_z - \frac{HHI_m^T}{s_m^T} \Delta ICE_m^{\mathrm{lin}} \right] = \Delta ICE_m - \Delta ICE_m^{\mathrm{lin}}$.

D Weighting: Summary of Estimates by Weighting Scheme

In my main market-level estimates I weight local labor markets by their 1991 formal employment, so that they describe the market of the average formal worker and are identified by the tariff cuts. In this appendix, I summarize what changes when each market instead counts once, so that the estimates describe the average market, and which weights I use to estimate each of the model's parameters.

Table D.2 summarizes. Weighting changes which comparison identifies my market-level effects, but not their sign: weighted, the effects on concentration and on the number of firms come from differences in tariff reductions across sectors and survive the recentering and permutation checks, while the effect on employment rests more on the contrast between markets with and without tradable firms; unweighted, the effects on concentration and employment come from the contrast between the 85% of markets with a tradable firm in 1991 and the 15% without one, and rest on parallel trends between the two groups, which hold for concentration and not for employment (unweighted, the effect on employment is 0.01, SE of 0.09, with the recentered measure, and −1.27, SE of 0.12, with the indicator for markets with a tradable firm; Table D.3). The two agree within quintiles of market size (Table D.1). Relative to the baseline level of concentration, the unweighted effect is the smaller (7% against 22%), because the markets where formal employment is located are much less concentrated than the average market. At the microregion level nearly every market has a tradable firm (6 of 475 do not), so the contrast is absent and I weight by 1991 formal employment, as in the extended model. The model's elasticities are not averages over markets, so I choose the weights in their estimation for identification and precision: I estimate $1/\eta$ with market fixed effects and markets weighted by 1991 employment (with each firm-market pair counting once, I estimate 1.226, SE of 0.160); I estimate $1/\theta$ unweighted, because with employment weights its first stage has no power; and I estimate the extended model's $1/\tilde{\rho}$, $1/\tilde{\theta}$ and Ω with the weights under which their first stages are strong (Appendix F.3). Because I estimate $1/\theta$ with each market counting once, I use the unweighted effect on concentration to compute the effect on markdowns in Table 4; with the weighted effect it is 0.011 instead of 0.013, a reduction in the take-home share of 0.30 instead of 0.37 cents.

What the unweighted effect on concentration captures. Unweighted, a 10% increase in exposure raises the payroll Herfindahl by 0.021 points, 7% of the 0.29 mean across markets (Appendix Table A.4, column (4)). This effect is identified by the shares: it is close to zero when I control for the expected value of ΔICE_m under random reassignment of tariff cuts across sectors (−0.005, SE of 0.021, per unit of ΔICE_m, against 0.212; Table D.3), and it falls to 0.002 points per 10% increase in exposure when I drop the markets without a tradable firm (Table D.4). Interpreting it as causal requires that markets with and without tradable firms would have followed the same trends absent the reform, which the flat 1986–1989 coefficients support (p-value of 0.35 on their trend; Table D.3). The comparison captures a smaller decline in concentration: between 1991 and

2000 the payroll Herfindahl fell by 0.15 points on average in the markets with no tradable firm and by 0.06 points in the markets with one. The gap is in non-tradable firms, whose contribution to the Herfindahl fell by 0.17 points in the markets with no tradable firm, where they are all of the firms, and by 0.02 points in the others.

Table D.1: Effect of import competition on concentration by quintile of 1991 market size

	Δ Payroll Herfindahl (premium-based)		Δ Employment Herfindahl	
	Unweighted	Weighted	Unweighted	Weighted
	(1)	(2)	(3)	(4)
Quintile 1 (smallest)	0.274	0.203	0.326	0.245
	(0.045)	(0.044)	(0.044)	(0.042)
Quintile 2	0.128	0.130	0.178	0.177
	(0.033)	(0.034)	(0.030)	(0.032)
Quintile 3	0.086	0.085	0.120	0.114
	(0.030)	(0.031)	(0.027)	(0.028)
Quintile 4	0.026	0.044	0.042	0.049
	(0.033)	(0.034)	(0.029)	(0.030)
Quintile 5 (largest)	0.148	0.173	0.120	0.097
	(0.033)	(0.049)	(0.029)	(0.027)

Notes: Standard errors in parentheses, clustered at the local labor market level. Each cell reports the 2000 coefficient of the stacked differences-in-differences regression of Table 1, estimated separately within each quintile of local labor markets' 1991 formal employment, unweighted (columns (1) and (3)) and weighted by 1991 formal employment (columns (2) and (4)).

Table D.2: Market-level estimates and parameters by weighting scheme

	Weighted by 1991 employment (average worker)	Unweighted (average market)	Identified by
Panel A: Market-level effects, per unit of ΔICE_m, 1991–2000			
Δ Payroll Herfindahl	0.172	0.212	Weighted: tariff cuts; unweighted: exposed vs. never exposed
	(0.042)	(0.017)	
Δ Average wage markdown, γ, per 10%	0.011	0.013	via β
Implied change in take-home share, cents per 10%	−0.30	−0.37	
Microregion level (extended model, ΔICE^*), per 10%			
Δ Expected payroll Herfindahl	0.028	–	
	(0.007)		
Δ Wage work share of market payroll	−0.032	–	
	(0.005)		
Δ Average inverse elasticity (take-home, cents)	0.0053 (−0.15)	–	tariff cuts (Table 7)
	(0.0014)		
Panel B: Parameters and the weights used in their estimation			
$1/\eta$ (firm × market; weights: market 1991 employment)	0.832 (0.093)	1.226 (0.160)	tariff cuts across firms within markets
$1/\theta$ (market; unweighted)	no first stage (F below 1)	1.454 (0.196)	exposed vs. never exposed (employment); tariff cuts (wage index)
$1/\tilde{\rho}$, $1/\tilde{\theta}$ (census cells; weights: 1991 formal employment)	0.684 (0.234), 1.187 (0.481)	–	ΔICE^* across cells
Ω (firm × cell; weights: lagged cell employment)	0.0091 (0.0025)	–	firm tariff × ΔICE^*

Notes: Panel A reports the 2000 coefficient of equation (12) on ΔICE_m, weighted by the market's 1991 formal employment and unweighted; standard errors clustered by local labor market. The microregion rows are the extended model's reduced forms on ΔICE^*, with the take-home share change in cents in parentheses. Panel B lists the sample, the weights and the variation that identifies each parameter; the extended model's parameters are in Table 7. Sources: Tables 1, 2, 3, 4, and 7, and Appendix Tables A.4, column (4), and D.3.

Table D.3: Effects by source of variation and pre-period tests

		ΔICE_m (1)	Recentered ΔICE_m (2)	Exposed indicator (3)	Pre-period trend (p-value) (4)
Δ Payroll Herfindahl (premium-based)	Weighted	0.172	0.151	0.464	0.016
		(0.042)	(0.036)	(0.062)	
	Unweighted	0.212	−0.005	0.590	0.346
		(0.017)	(0.021)	(0.033)	
Δ Log number of firms	Weighted	−0.971	−1.021	−0.754	0.000
		(0.188)	(0.214)	(0.296)	
	Unweighted	−0.510	−0.300	−0.873	0.469
		(0.048)	(0.063)	(0.085)	
Δ Log formal employment	Weighted	−0.398	−0.242	−0.625	0.000
		(0.186)	(0.206)	(0.213)	
	Unweighted	−0.401	0.010	−1.266	0.027
		(0.069)	(0.090)	(0.123)	
Δ Log formal wage premium	Weighted	−0.105	−0.112	−0.186	0.308
		(0.089)	(0.082)	(0.109)	
	Unweighted	0.027	0.170	−0.343	0.004
		(0.030)	(0.038)	(0.060)	

Notes: Coefficient on ΔICE_m for the change between 1991 and 2000 in equation (12). Column (2) adds as a control, interacted with year, the expected value of ΔICE_m when tariff changes are reassigned at random across CNAE sectors within IBGE subsectors (Borusyak and Hull, 2023). Column (3) replaces ΔICE_m with an indicator for exposed markets (markets with firms facing tariff cuts in 1991). Column (4): p-value of the slope of a linear trend fitted to the 1986–1989 coefficients of column (1). Standard errors clustered by local labor market in parentheses.

Table D.4: Controls for agriculture intensity and tradable sector intensity, unweighted

	Δ Payroll Herfindahl (premium-based) (1)	Δ Employment Herfindahl (2)	Δ Log formal employment (3)
Main specification, unweighted	0.212	0.239	-0.401
	(0.017)	(0.016)	(0.069)
Control: 1991 census agricultural share × year	0.174	0.201	-0.254
	(0.017)	(0.016)	(0.069)
ΔICE without agricultural firms	0.216	0.244	-0.342
	(0.017)	(0.016)	(0.068)
Dropping top-quartile agricultural microregions	0.176	0.195	-0.306
	(0.018)	(0.018)	(0.074)
Control: Dix-Carneiro and Kovak (2017)'s RTR × year	0.165	0.191	-0.219
	(0.017)	(0.016)	(0.069)
Control: 1991 payroll share of tariff firms × year	0.230	0.255	-0.338
	(0.017)	(0.016)	(0.069)
Control: 1{tariff firms in 1991} × year	0.020	0.016	0.039
	(0.018)	(0.017)	(0.080)
Markets with tariff firms only	0.020	0.016	0.039
	(0.018)	(0.017)	(0.080)
Microregion × year fixed effects	0.163	0.190	-0.041
	(0.019)	(0.018)	(0.075)
Occupation × year and size decile × year FE	0.129	0.151	0.270
	(0.018)	(0.017)	(0.065)

Notes: Standard errors in parentheses, clustered at the local labor market level. All entries report the 2000 coefficient on ΔICE_m from the stacked differences-in-differences regressions of Table 1, with each local labor market counting once. The agricultural employment share is the microregion's 1991 share of tradable employment in agriculture, from the census. The RTR is Dix-Carneiro and Kovak (2017)'s RTR exposure measure, computed from the 1991 census. The tariff firms' payroll share is the 1991 payroll share of the market's firms in sectors with a tariff. ΔICE_m without agricultural firms excludes agricultural firms from the market's tariff firms. The top quartile of agricultural share is defined across microregions.

E Effective Rates of Protection

The tariff shocks of the paper are changes in output tariffs. Brazil's liberalization also cut tariffs on firms' inputs, which lowers their costs and works against the import competition they face. In this appendix, I repeat the market-level estimates of Table 1 and the two elasticities of Tables 2 and 3 with the shock measured as the change in effective rates of protection (ERP), output tariffs net of input tariffs using Brazil's 1985 input-output table (Appendix A). Input tariffs are available for about 50 sectors, against 285 for output tariffs, so the ERP shock is noisier and its effects attenuated; the ERP version of ΔICE_m also weights firms by their linear payroll shares rather than by their contributions to the payroll Herfindahl. In Table E.1 I therefore report, for each weighting of markets, the output-tariff shock of Table 1, the output-tariff shock with linear payroll shares, and the ERP shock with the same shares, so that I can tell the two differences apart.

Weighted by 1991 formal employment (columns (1)–(3)), the effect on concentration is the same with the ERP shock as in the main specification (21% against 22% of the 1991 mean), the number of firms falls by less, formal employment does not fall, and the three wage measures rise by 0.16 to 0.18 log points per unit of the shock (SE of 0.06 to 0.08). Column (2) shows that part of the difference comes from the share weights rather than from input tariffs: with output tariffs and linear payroll shares, the wage effects are also positive, though imprecise (0.20 to 0.26, SE of 0.11 to 0.15). Unweighted (columns (4)–(6)), the ERP shock gives effects on concentration, the number of firms and employment of the same sign as the output-tariff shock and about half as large relative to the 1991 mean (4% against 7% for the payroll Herfindahl), and a positive effect on the wage premium (0.065, SE of 0.020), where the output-tariff shock gives zero. The within-market elasticity is nearly unchanged with the ERP shock (0.878 against 0.832; Table E.2), with first stage and reduced form both about a third smaller, and the cross-market inverse elasticity is larger and less precise (1.758, SE of 0.421, against 1.454), so that the gap between them and the implied effect on markdowns are larger. Because the positive wage effects of the ERP shock partly reflect the weighting of firms within the exposure measure, and because input tariffs are measured at a much coarser level than output tariffs, I keep output tariffs as the shock in the main text and leave the separate roles of input and output tariff reductions for future work.

Table E.1: Effect of import competition on market outcomes: effective rates of protection

	Weighted by 1991 formal employment			Unweighted		
	Main specification	Linear payroll shares	Effective rates of protection	Main specification	Linear payroll shares	Effective rates of protection
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Formal sector labor market concentration						
Δ Payroll Herfindahl (premium-based)	0.172	0.421	0.164	0.212	0.262	0.121
	(0.042)	(0.074)	(0.039)	(0.017)	(0.020)	(0.011)
1991 mean	0.079	0.079	0.079	0.287	0.287	0.287
Effect per 10%, percent of 1991 mean	22%	54%	21%	7%	9%	4%
Δ Payroll Herfindahl	0.183	0.439	0.173	0.213	0.262	0.122
	(0.045)	(0.083)	(0.043)	(0.017)	(0.020)	(0.011)
1991 mean	0.082	0.082	0.082	0.292	0.292	0.292
Effect per 10%, percent of 1991 mean	22%	53%	21%	7%	9%	4%
Δ Employment Herfindahl	0.109	0.278	0.101	0.239	0.294	0.134
	(0.024)	(0.035)	(0.019)	(0.016)	(0.019)	(0.011)
1991 mean	0.056	0.056	0.056	0.239	0.239	0.239
Effect per 10%, percent of 1991 mean	19%	50%	18%	10%	12%	6%
Panel B: Number of firms and formal employment						
Δ Log number of firms	−0.971	−1.443	−0.307	−0.510	−0.625	−0.277
	(0.188)	(0.196)	(0.111)	(0.048)	(0.054)	(0.032)
Δ Log formal employment	−0.398	−0.048	0.128	−0.401	−0.479	−0.194
	(0.186)	(0.199)	(0.100)	(0.069)	(0.078)	(0.046)
Panel C: Wages						
Δ Log average firm wage premium	−0.124	0.200	0.157	−0.071	−0.085	0.022
	(0.130)	(0.151)	(0.075)	(0.035)	(0.040)	(0.024)
Δ Log formal wage premium	−0.105	0.208	0.180	0.027	0.039	0.065
	(0.089)	(0.113)	(0.056)	(0.030)	(0.035)	(0.020)
Δ Log formal average wage	−0.075	0.255	0.183	0.014	0.020	0.067
	(0.110)	(0.131)	(0.065)	(0.035)	(0.040)	(0.023)
Observations	296,311	296,311	296,311	296,311	296,311	296,311
Local labor markets	19,757	19,757	19,757	19,757	19,757	19,757

Notes: Standard errors in parentheses, clustered at the local labor market level. Each cell reports the 2000 coefficient from a stacked differences-in-differences regression of long-differenced outcomes on the trade shock interacted with year dummies, with local labor market fixed effects, as in Table 1. Columns (1) and (4) repeat Table 1, with ΔICE_m built from output tariffs with weights given by firms' contributions to the market's 1991 payroll Herfindahl (Table 1, column (1), and Appendix Table A.4, column (4)); columns (2) and (5) build ΔICE_m from output tariffs with firms' linear 1991 payroll shares as weights (Appendix Table A.4, column (2)); columns (3) and (6) build it from effective rates of protection with linear payroll shares. Columns (1)–(3) weight local labor markets by their 1991 formal employment; columns (4)–(6) count each local labor market once. Effective rates of protection are the output tariff net of the input-output-weighted average of input tariffs, over one minus the input share, computed with Brazil's 1985 input-output table at the Nível 50 sector level (Appendix A).

Table E.2: Labor supply elasticities with effective rates of protection

	$1/\eta$ (Table 2)		$1/\theta$ (Table 3)	
	Import tariffs (1)	Effective rates of protection (2)	Import tariffs (3)	Effective rates of protection (4)
Panel A: First stage				
Δ Firm log employment in LLM	−0.736	−0.447		
	(0.076)	(0.059)		
Δ LLM employment index			−0.294	−0.114
			(0.044)	(0.029)
First stage F	94.9	57.4	45.6	15.7
Panel B: Reduced form				
Δ Firm wage premium in LLM	−0.612	−0.392		
	(0.036)	(0.029)		
Δ LLM wage premium index			−0.183	−0.100
			(0.050)	(0.035)
Panel C: 2SLS				
$1/\eta$	0.832	0.878		
	(0.093)	(0.124)		
$1/\theta - 1/\eta$			0.622	0.880
			(0.217)	(0.438)
$1/\theta$			1.454	1.758
			(0.196)	(0.421)
Implied bound on wage take-home share	55%	53%	41%	36%
Observations	363,407	362,392	17,628	17,628
Firms	149,083	148,857		
Local labor markets	11,632	11,603	17,628	17,628

Notes: Columns (1) and (3) reproduce column (1) of Tables 2 and 3. Columns (2) and (4) replace the 1990–1994 change in the firm's import tariff by the change in its effective rate of protection as the instrument, in the firm-level regressions of Section 6.1 (standard errors clustered by firm; local labor market fixed effects; weighted by baseline market employment; 1986–1991 pre-trend controls) and in the market-level regressions of Section 6.2 (standard errors clustered by local labor market; state fixed effects and 1986–1990 trends), with ΔICE_m in column (4) built from effective rates of protection as in Appendix Table E.1, column (6). Column (4) uses the $1/\eta$ of column (2). The implied bound on the take-home share is $1/(1+1/\eta)$ in columns (1)–(2), an upper bound, and $1/(1+1/\theta)$ in columns (3)–(4), a lower bound.

F Extension: Oligopsony in Dual Labor Markets

F.1 Model

I model the labor supply preferences of *people seeking formal employment*. Firms exploit these preferences to set wages below workers' marginal revenue product when maximizing profits while holding local competitors' labor demand decisions constant. The starting point is the labor supply preference structure of Section 2.1, which I extend in two directions. First, I split each local labor market into two sectors, wage work and self-employment. Second, I allow workers to take into account the probability of involuntary separation out of the formal sector—which varies by firm—when making formal sector labor supply decisions.

F.1.1 Discrete choice formal labor supply in dual labor markets

There is a continuum of homogeneous workers j, each choosing where to supply l^j effective units of labor.[60] Each worker chooses a single market m from a large but finite number of local labor markets. Within each market, workers choose a sector g, which is either wage work $\overline{g}$ or self-employment $\underline{g}$. When considering wage work, workers compare expected earnings at a finite number of formal sector firms z, which reflect their perceived probability of involuntary separation out of the formal sector o in market m. When considering self-employment, they consider expected earnings from starting a small business in market m. Workers choose the employment option that minimizes their disutility of labor given their preference parameters, expected earnings, and a minimum earnings requirement, as in Section 2.1.

Two features distinguish the extension. First, workers value formal jobs at *expected* wages. A job at formal firm z carries a perceived probability $p_{z\overline{g}m}$ of involuntary separation from formal employment. I assume that a separated worker finds informal wage work without frictions, in a perfectly competitive local labor market for informal labor that pays w_m^o. The wage relevant for attracting labor to the firm is therefore

$$\bar{w}_{z\overline{g}m} = \left(1 - p_{z\overline{g}m}\right) w_{z\overline{g}m} + p_{z\overline{g}m}\, w_m^o = w_{z\overline{g}m}\left(1 + p_{z\overline{g}m}\, \sigma_{z\overline{g}m}\right), \tag{F.1}$$

where $w_{z\overline{g}m}$ is the firm's posted formal wage and $\sigma_{z\overline{g}m} \equiv (w_m^o - w_{z\overline{g}m})/w_{z\overline{g}m}$ is the wage gap between the local informal sector and the firm's formal wage. This gap may be negative (the informal sector pays less) or positive (the informal sector pays more). Workers likewise value self-employment at its expected returns $W_{\underline{g}m}$, since the returns to starting a small business are uncertain ex ante. Introducing earnings uncertainty is a natural way to incorporate both self-employment and informal

[60] As in the baseline model, the assumption of homogeneous productivity is mapped to the data by conditioning wages on worker demographics.

wage work into the labor supply decisions of people seeking formal sector employment.

No dynamic gains from influencing workers' priors. I assume that firms do not internalize that their hiring and firing decisions at time t influence workers' priors about the probability of separation from the firm at $t+1$ or after. The assumption shuts off $p_{z\overline{g}m}$ as an additional, dynamic lever of wage setting, and thereby preserves the baseline model's static wage-setting problem while letting priors enter labor supply. Workers may form priors from firms' past involuntary separations, which the administrative data record, but neither workers nor firms weigh the future costs or benefits of decisions made at t. Dynamic gains to labor market power from firms' ability to shape workers' priors about separation, and about other within-firm dynamics such as promotion and tenure, merit further research.

Second, idiosyncratic tastes follow a nested GEV distribution with a sector level between the firm and market levels of the baseline model:

$$G(\xi) = \exp\left(-\sum_m \left[\sum_{g\in\{\overline{g},\underline{g}\}} \left(\sum_z e^{-(1+\tilde{\eta})\,\xi^j_{zgm}}\right)^{\frac{1+\tilde{\rho}}{1+\tilde{\eta}}}\right]^{\frac{1+\tilde{\theta}}{1+\tilde{\rho}}}\right). \tag{F.2}$$

The parameters $\tilde{\eta}$, $\tilde{\rho}$, and $\tilde{\theta}$ are workers' elasticities of substitution across firms within the wage work sector, between wage work and self-employment within markets, and across markets, respectively. $\tilde{\eta}$ and $\tilde{\theta}$ are the extended counterparts of the baseline model's η and θ, and all three elasticities are common to all workers.

F.1.2 Labor supply curves faced by individual formal firms

As in the baseline model, the discrete choice formulation of workers' labor supply decisions with GEV idiosyncratic tastes yields analytic expressions for the probability that worker j chooses formal sector firm z in the wage work sector $\overline{g}$ of market m. Integrating these probabilities over the continuum of workers gives total labor supply to formal sector firm z:

$$l_{z\overline{g}m} = L\left(\frac{\bar{w}_{z\overline{g}m}}{W_{\overline{g}m}}\right)^{\tilde{\eta}}\left(\frac{W_{\overline{g}m}}{W_m}\right)^{\tilde{\rho}}\left(\frac{W_m}{W}\right)^{\tilde{\theta}} \times \left[\xi^{1+\tilde{\eta}}_{zgm}\xi^{1+\tilde{\rho}}_{gm}\xi^{1+\tilde{\theta}}_{m}\right]^{-1}. \tag{F.3}$$

whose inverse is:

$$\bar{w}_{z\overline{g}m} = W\left(\frac{l_{z\overline{g}m}}{L_{\overline{g}m}}\right)^{\frac{1}{\tilde{\eta}}}\left(\frac{L_{\overline{g}m}}{L_m}\right)^{\frac{1}{\tilde{\rho}}}\left(\frac{L_m}{L}\right)^{\frac{1}{\tilde{\theta}}} \xi^{1+\frac{1}{\tilde{\eta}}}_{z\overline{g}m}\xi^{1+\frac{1}{\tilde{\rho}}}_{\overline{g}m}\xi^{1+\frac{1}{\tilde{\theta}}}_{m}. \tag{F.4}$$

Here $(W_{\overline{g}m}, L_{\overline{g}m})$, (W_m, L_m), and (W, L) are nested CES wage and labor supply indices over firms, sectors, and markets, defined at the end of this subsection. Equation F.4 has the same form as its

baseline counterpart: the wage needed to attract labor is increasing in the firm's employment relative to its sector, in the sector's employment relative to the market, and in the market's employment relative to the country. The difference is that the relevant wage is the expected wage of equation F.1, which depends on the probability of separation and on local informal sector conditions.

In addition, because I can only estimate the elasticity of substitution between self-employment and wage work with census data—which do not include firm boundaries but do include regions—I introduce the market-level aggregate of equation F.4, the inverse labor supply curve to wage work in each market:

$$W_{\bar{g}m} = W \left(\frac{L_{\bar{g}m}}{L_m} \right)^{\frac{1}{\bar{\rho}}} \left(\frac{L_m}{L} \right)^{\frac{1}{\bar{\theta}}} \xi_{\bar{g}m}^{1+\frac{1}{\bar{\rho}}} \xi_m^{1+\frac{1}{\bar{\theta}}}. \tag{F.5}$$

Finally, the wage and labor supply indices in equations F.3–F.5 are nested CES:

$$W_{\bar{g}m} = \left[\sum_z \left(\frac{\bar{w}_{z\bar{g}m}}{\xi_{z\bar{g}m}} \right)^{1+\tilde{\eta}} \right]^{\frac{1}{1+\tilde{\eta}}}, \qquad L_{\bar{g}m} = \left[\sum_z \left(\xi_{z\bar{g}m} l_{z\bar{g}m} \right)^{\frac{1+\tilde{\eta}}{\tilde{\eta}}} \right]^{\frac{\tilde{\eta}}{1+\tilde{\eta}}} \tag{F.6}$$

$$W_m = \left[\left(\frac{W_{\underline{g}m}}{\xi_{\underline{g}m}} \right)^{1+\tilde{\rho}} + \left(\frac{W_{\bar{g}m}}{\xi_{\bar{g}m}} \right)^{1+\tilde{\rho}} \right]^{\frac{1}{1+\tilde{\rho}}}, \qquad L_m = \left[\left(\xi_{\underline{g}m} l_{\underline{g}m} \right)^{\frac{1+\tilde{\rho}}{\tilde{\rho}}} + \left(\xi_{\bar{g}m} L_{\bar{g}m} \right)^{\frac{1+\tilde{\rho}}{\tilde{\rho}}} \right]^{\frac{\tilde{\rho}}{1+\tilde{\rho}}} \tag{F.7}$$

$$W = \left[\sum_m \left(\frac{W_m}{\xi_m} \right)^{1+\tilde{\theta}} \right]^{\frac{1}{1+\tilde{\theta}}}, \qquad L = \left[\sum_m \left(\xi_m L_m \right)^{\frac{1+\tilde{\theta}}{\tilde{\theta}}} \right]^{\frac{\tilde{\theta}}{1+\tilde{\theta}}} \tag{F.8}$$

where $W_{\underline{g}m}$ are expected earnings from self-employment in market m.[61]

F.1.3 Firm-specific wage markdowns

Firms compete for labor à la Cournot, as in Section 2.2. Each firm chooses employment while holding constant competing firms' labor demand, workers' priors about separation probabilities, expected informal sector wages, and expected returns to self-employment. Taking logs of equation F.4 and differentiating with respect to the firm's own employment gives the inverse elasticity of

[61] In a nested CES preference structure, $W_{\underline{g}m}$ equals expected log earnings in the limit where the elasticity of substitution between self-employment options is zero. At the micro level, this means that no one worker can provide self-employment for another worker. At the macro level, this is the CES limiting case of Cobb-Douglas preferences over self-employment options. Formally, let $\eta_{\underline{g}}$ denote the elasticity of substitution between self-employment options and let $\mathbb{E}[\pi_{\underline{g}m}^j]$ denote worker j's expected earnings from self-employment in market m. Then $W_{\underline{g}m} = \left(\int_0^1 \mathbb{E}[\pi_{\underline{g}m}^j]^{1+\eta_{\underline{g}}} \, dj \right)^{1/(1+\eta_{\underline{g}})}$ is the CES aggregator over self-employment options in a given market. When $\eta_{\underline{g}} \to 0$, we have $W_{\underline{g}m} \to \exp\left(\int_0^1 \log \mathbb{E}[\pi_{\underline{g}m}^j] \, dj \right)$, the Cobb-Douglas form. Note that $W_{\underline{g}m} > 0$ because workers must satisfy their minimum earnings requirement in equilibrium.

residual labor supply that the firm faces:

$$\varepsilon_{z\bar{g}m}^{-1} = \left[\frac{1}{\tilde{\rho}} s_{z\bar{g}m} + \frac{1}{\tilde{\eta}}(1 - s_{z\bar{g}m})\right] + \left(\frac{1}{\tilde{\theta}} - \frac{1}{\tilde{\rho}}\right) s_{zm}. \tag{F.9}$$

Here $s_{z\bar{g}m}$ is firm z's share of the wage work sector's expected payroll and s_{zm} is its share of the market's total expected payroll, including self-employment. The elasticity of labor supply to a formal firm thus depends on how easily workers substitute between firms within markets ($\tilde{\eta}$), between markets ($\tilde{\theta}$), and between wage work and self-employment within markets ($\tilde{\rho}$), and the relevant shares are based on expected wages.

Local average wage markdown. As in the baseline model, the average wage markdown is $\mu_{\bar{g}m} \equiv 1 + \varepsilon_{\bar{g}m}^{-1}$ and the wage take-home share is $\mu_{\bar{g}m}^{-1}$. Taking a weighted average of the firm-specific inverse elasticities in equation F.9, with $s_{z\bar{g}m}$ as weights, gives the market's average inverse elasticity of residual labor supply:

$$\bar{\varepsilon}_{\bar{g}m}^{-1} = \frac{1}{\tilde{\rho}} \widetilde{HHI}_{\bar{g}m} + \frac{1}{\tilde{\eta}}(1 - \widetilde{HHI}_{\bar{g}m}) + \left(\frac{1}{\tilde{\theta}} - \frac{1}{\tilde{\rho}}\right) \widetilde{HHI}_{\bar{g}m} \cdot s_m. \tag{F.10}$$

Here $\widetilde{HHI}_{\bar{g}m} \equiv \sum_z s_{z\bar{g}m}^2$ is the wage work sector's expected payroll Herfindahl and s_m is the wage work sector's share of market m's expected payroll.

F.1.4 Policy effects on wage markdowns

Let X_m denote an exogenous shock to labor demand in local labor market m. Then this shock's effect on local markdowns is given by its effect on average local inverse elasticities of labor supply, which follows:

$$\frac{\partial \bar{\varepsilon}_{\bar{g}m}^{-1}}{\partial X_m} = \left(\frac{1}{\tilde{\rho}} - \frac{1}{\tilde{\eta}}\right) \times \beta_{\text{HHI}} + \left(\frac{1}{\tilde{\theta}} - \frac{1}{\tilde{\rho}}\right) \left[s_m \times \beta_{\text{HHI}} + \widetilde{HHI} \times \beta_{s_m}\right] \tag{F.11}$$

where $\widetilde{HHI}$ is short for $\widetilde{HHI}_{\bar{g}m}$, and β_{HHI} and β_{s_m} are the reduced-form effects of the shock on the expected payroll Herfindahl and on the wage work share. Without self-employment ($s_m = 1$ and $\beta_{s_m} = 0$), the expression reduces to Corollary 1. In the first term, concentration shifts weight between the cross-firm margin $\tilde{\eta}$ and the self-employment margin $\tilde{\rho}$. The second term arises because the weight of the cross-market margin in equation F.10 is $\widetilde{HHI}_{\bar{g}m} \times s_m$, which rises with concentration and falls when the wage work share falls.

F.2 Trade shock

The baseline measure of import competition exposure, ΔICE_m (equation 11), averages the tariff reductions faced by the formal firms of a market. In the extended model, the market also includes informal wage work and self-employment. For the extended model, I use a measure with two components, weighted by the formal share of the market's wage bill.

The first component is ΔICE_m of equation 11, computed with microregions as markets: the average tariff reduction faced by the microregion's tradable sector firms, weighted by their 1991 squared payroll shares as in equation 11. It is zero in the 6 of the 475 microregions of the sample without a tradable sector firm with a tariff match, which remain in all estimations.[62]

The second component is RTR_m, Dix-Carneiro and Kovak (2017)'s exposure measure, RTR, with industry weights computed from the 1991 census employment of all workers in tradable industries, as in Kovak (2013), where the return to labor in a regional market moves with the industry composition of all of the region's workers.

The extended model's shock weights the two components by the formal share of the market's 1991 wage bill:

$$\Delta ICE^*_m \equiv f_{m,1991}\, \Delta ICE_m + \left(1 - f_{m,1991}\right) RTR_m, \tag{F.12}$$

where I measure $f_{m,1991}$ in the 1991 census as the earnings of formal employees over the earnings of all employed workers: formal employees, informal employees, and the self-employed. The shock is defined for each microregion and assigned to all demographic cells of the microregion. Across microregions, ΔICE^*_m has a mean of 0.084 and a standard deviation of 0.037.

F.3 Estimation strategy

I define local labor markets as microregions for all of the extension's estimates.[63] Unless noted, Δ denotes 1991–2000 long differences and standard errors are clustered by microregion.

F.3.1 Substitution within markets, across firms: $1/\eta$

The labor supply response to firms' posted wages is the baseline model's within-market cross-firm inverse elasticity. I estimate it as in Section 5, with markets defined as microregions:

$$\Delta \ln w_{zm} = \frac{1}{\eta} \Delta \ln l_{zm} + \delta_m + \epsilon_{zm}, \tag{F.13}$$

[62] Appendix C.2 relates the squared-share weights to payroll-share weights and to the baseline model's measure of exposure.

[63] While the 1991 and 2000 censuses contain occupation codes, they are missing for a large share of observations, and the two years also follow different definitions from each other (and from RAIS).

where Δ denotes 1991–2000 long differences, w_{zm} is firm z's December wage premium in microregion m, estimated on all of the firm's workers in the microregion as in the baseline, l_{zm} is its employment, and δ_m is a microregion fixed effect. I instrument the change in employment with the firm-level tariff reduction. Firms in non-tradable sectors remain in the sample with a zero tariff change. I weight the regression by markets' 1991 employment and cluster standard errors by firm.

F.3.2 The bias from involuntary separation out of the formal sector: Ω and $1/\tilde{\eta}$

If the extended model is true, workers respond to expected rather than posted wages, and the IV estimate of equation F.13 recovers only part of the relevant elasticity. Taking logs of equation F.1 gives

$$\ln \bar{w}_{z\bar{g}m} = \underbrace{\ln w_{z\bar{g}m}}_{\text{Formal wage}} + \underbrace{\ln\left(1 + p_{z\bar{g}m}\sigma_{z\bar{g}m}\right)}_{\text{Expected wage gap}} \tag{F.14}$$

and taking logs of equation F.4, with δ_m denoting a market fixed effect, gives

$$\ln \bar{w}_{z\bar{g}m} = \frac{1}{\tilde{\eta}} \ln l_{z\bar{g}m} + \delta_m + \ln \xi_{z\bar{g}m}^{1+\tilde{\eta}} \tag{F.15}$$

Bias formula. Given equation F.15, $1/\tilde{\eta}$ is the second-stage coefficient from an IV regression in which $\ln l_{z\bar{g}m}$ is instrumented by an exogenous labor demand shock $X_{z\bar{g}m}$. Expanding expected wages per equation F.14, and letting $(*)$ denote partialled-out variables to apply the Frisch-Waugh-Lovell theorem, gives

$$\frac{1}{\tilde{\eta}} = \frac{\text{Cov}(\ln \bar{w}^*_{z\bar{g}m}, X_{z\bar{g}m})}{\text{Cov}(\ln l^*_{z\bar{g}m}, X_{z\bar{g}m})} = \underbrace{\frac{\text{Cov}(\ln w^*_{z\bar{g}m}, X_{z\bar{g}m})}{\text{Cov}(\ln l^*_{z\bar{g}m}, X_{z\bar{g}m})}}_{\frac{1}{\eta}\text{: Supply response to formal wage}} + \underbrace{\frac{\text{Cov}[\ln(1 + p_{z\bar{g}m}\sigma_{z\bar{g}m})^*, X_{z\bar{g}m}]}{\text{Cov}(\ln l^*_{z\bar{g}m}, X_{z\bar{g}m})}}_{\Omega\text{: Supply response to expected wage gap}} \tag{F.16}$$

The first term in equation F.16 is the response to formal wages, $1/\eta$, which equation F.13 estimates. The second term, Ω, is the labor supply response to changes in expected formal–informal wage gaps, conditional on changes in formal wages. It is the bias in the baseline model's elasticity if the extended model is true. If workers do not take the possibility of separation into account, the baseline model holds and $1/\eta$ is the relevant elasticity. The sign of Ω depends on local informal wages and on the joint distribution, across the formal firms of a market, of separation probabilities and formal wages.

Interpretation. If Ω is *negative*, *less* labor is supplied to firms with higher expected wage gaps relative to the same market's informal sector wage. That is, workers dislike firms that in expectation pay less than the local informal sector.[64] Conversely, if Ω is *positive*, *more* labor is

[64] The possibility that informal sector options might pay more than formal sector options should not be discarded.

supplied to firms with higher expected wage gaps relative to the same market's informal sector wage. In this case, the inverse elasticity of substitution $1/\eta$ estimated based on the baseline model is *too small* relative to the true inverse elasticity $1/\tilde{\eta}$, meaning that *firms have more labor market power* than originally estimated.

Regression. I estimate Ω from the second-stage regression

$$\Delta \ln \left(1 + p_{z\bar{g}m}\, \sigma^{c}_{z\bar{g}m}\right) = \Omega\, \Delta \ln l^{c}_{z\bar{g}m} + \delta_z + \delta_m + \delta_c + \epsilon^{c}_{z\bar{g}m}, \tag{F.17}$$

where an observation is a firm × microregion × demographic cell, the left-hand side is the 1991–2000 change in the expected wage gap of cell c at firm z, the regressor is the change in the firm's log employment of that cell, and the δ terms are firm, microregion, and cell fixed effects. Including a firm fixed effect ensures that Ω captures the supply response to changes in expected wage gaps relative to the informal sector conditional on firm-level changes in wage levels, the identifying variation for η in both the baseline and extended models. I weight observations by the lagged log of the firm's employment of the cell and cluster standard errors by microregion.

Measurement. RAIS records the reason for every separation, distinguishing firings from quits. I proxy workers' prior about the probability of involuntary separation from firm z in microregion m with the firm's fired-worker share,

$$p_{z\bar{g}m} = \frac{\text{fired}_{zm}}{\text{fired}_{zm} + \text{stayers}_{zm}}, \tag{F.18}$$

where fired_{zm} is the number of workers fired from firm z in microregion m during the calendar year and stayers_{zm} is the firm's employment in the microregion on December 31. The wage gap $\sigma^{c}_{z\bar{g}m}$ compares the informal monthly earnings of cell c in microregion m, measured in the census, with the cell's December earnings at firm z, measured in RAIS and scaled by the ratio of mean census to mean RAIS formal earnings in each year. Earnings are in constant 2000 reais. I drop firm × cell earnings above the 99th percentile and wage gaps outside the 1st–99th percentile range. I merge RAIS and census data by microregion, demographic cell, and year.

Panel A of Figure F.1 plots a binned scatter of firm-level probabilities of involuntary separation against informal–formal wage gaps. The probability is on average 3%. It is highest in local labor markets where the informal sector pays less than the formal sector (for the same demographic cell). This means that workers who are at higher risk of separation would also face lower-paying informal sector jobs. Panel B of Figure F.1 shows that a substantial number of jobs are in local labor markets where the informal sector pays less, but there is a sizable number of jobs for which informal wage work pays higher wages than the formal sector. This is consistent with the residual

Appendix Figure F.1 shows that a non-trivial share of regions exhibit larger monthly earnings for workers of the same demographic composition than monthly earnings at formal sector firms.

wage distributions plotted in Figure F.2.

Instruments. Identification requires labor demand shocks that vary across firms within markets, as for η, and that shift expected wage gaps conditional on firm-level changes in formal wages. I instrument $\Delta \ln l^c_{z\bar{g}m}$ with the interaction between the firm-level tariff reduction and the microregion's trade shock,

$$X^c_{z\bar{g}m} = \left(1 + \Delta \ln \left(1 + \tau_{i(z)}\right)\right) \times \Delta ICE^*_m, \tag{F.19}$$

where $\Delta \ln(1 + \tau_{i(z)})$ is the firm-level tariff reduction of equation 17. I interact the instrument with gender, education, and age group indicators. The firm-level tariff reduction changes the firm's labor demand and workers' perceived probability of separation from it. The microregion's shock changes equilibrium informal wages. Because the regression includes firm and microregion fixed effects, the instrument's variation comes from firms that operate in several microregions, through differences in ΔICE^*_m across a firm's microregions scaled by its own tariff reduction, and from the firm's differential exposure across demographic cells through the interactions. In the first stage, the common component has coefficients of -13.5 for men and -13.4 for women, and the education and age differentials are small. The exclusion restriction is that, conditional on firm, microregion, and cell fixed effects, tariff reductions affect labor supply only through separation probabilities and equilibrium wages, and not through workers' tastes for individual firms. The standard error of $1/\tilde{\eta} = 1/\eta + \Omega$ combines those of its two components as $\sqrt{\mathrm{SE}(1/\eta)^2 + \mathrm{SE}(\Omega)^2}$, treating the two estimates as independent. Since both use the same firms and tariff reductions, they need not be independent; under any correlation between them, the standard error is at most $\mathrm{SE}(1/\eta) + \mathrm{SE}(\Omega) = 0.100$, against 0.097 under independence.

Unemployment insurance. Finally, I consider an alternative measure of the expected wage in which a separated worker receives unemployment insurance before moving to informal wage work. I value earnings after separation at the formal wage for four of the twelve months, the typical benefit at the time, and at the local informal wage for the remaining eight, so that the wage gap is two thirds of its unadjusted value.[65]

F.3.3 Substitution to self-employment and across markets

To arrive at the regression equation to estimate the elasticity of substitution between wage work and self-employment, take logs of equation F.5, express it in long-differences, add a constant to absorb

[65] Brazil's unemployment insurance was instituted by the 1990 law in Brasil (1990). Its rules in the 1990s were involved, with earnings brackets limiting benefit amounts and eligibility contingent on proof of time of service, but the typical benefit was four months of salary, paid monthly. Benefit brackets and their changes over time are documented at https://www.debit.com.br/tabelas/seguro-desemprego.

country-level changes, and group taste shifters into an error term $\epsilon_{\overline{g}m}$ to get:

$$\Delta \ln W_{\overline{g}m} = \alpha + \frac{1}{\tilde{\rho}} \Delta \ln \left(\frac{L_{\overline{g}m}}{L_m} \right) + \frac{1}{\tilde{\theta}} \Delta \ln \left(\frac{L_m}{L} \right) + \epsilon_{\overline{g}m} \tag{F.20}$$

A challenge in estimating equation F.20 is that each labor market has a single wage work sector, and import tariff reductions were primarily a labor demand shock to wage work. As a result, I cannot separately identify $\tilde{\rho}$ from $\tilde{\theta}$ unless I partition markets into groups, each with its own wage work and self-employment sectors, that have differential shocks to wage work labor demand (the within-market, cross-sector variation that identifies $\tilde{\rho}$) but are subject to the same market-level shock (the cross-market variation that identifies $\tilde{\theta}$).[66]

Regression. I partition labor markets into demographic cells, each with its own wage work and self-employment sectors. Let c denote the 18 cells defined by gender (men; women), education (primary; secondary; tertiary), and age (18–29; 30–49; 50–64). I estimate

$$\Delta \ln W^c_{\overline{g}m} = \frac{1}{\tilde{\rho}} \Delta \ln \left(\frac{L^c_{\overline{g}m}}{L^c_m} \right) + \frac{1}{\tilde{\theta}} \Delta \ln \left(\frac{L_m}{L} \right) + \delta_r + \epsilon^c_{\overline{g}m}, \tag{F.21}$$

where Δ denotes long differences within microregion $\times$ cell pairs and δ_r is a fixed effect for each of Brazil's five major regions, which absorbs regional general equilibrium effects of trade liberalization. The cells of a market share the market's employment share L_m/L, so the cross-market margin is common to them. The partition provides the within-market, cross-sector variation that identifies $1/\tilde{\rho}$.

Measurement. I estimate equation F.21 using Brazil's 1991 and 2000 censuses. The indices $W^c_{\overline{g}m}$, $L^c_{\overline{g}m}$, L^c_m, and L_m reflect lower-nest elasticities and taste shifters. To keep the analysis executable with census data, I use their directly observable counterparts. I measure $L^c_{\overline{g}m}/L^c_m$ as the wage work share of cell c's employment in microregion m, L_m/L as the microregion's share of national employment, and $\ln W^c_{\overline{g}m}$ as the mean log residual real monthly earnings of the cell's wage workers. Wage workers are formal and informal employees. I residualize earnings at the individual level on a fully saturated vector of gender, education, age, and year indicators, so that neither demographic composition nor demographic-specific wage trends enter the identifying variation. My sample consists of microregion $\times$ cell pairs with positive 1991 formal employment, and I weight the regression by the cell's 1991 formal employment.

Instruments. I instrument the two endogenous variables in equation F.21 with the microregion's trade shock ΔICE^*_m and its interactions with (a) age group, gender, and education indicators and (b) the cell's 1991 formal employment share. The shock is common to the cells of a microregion. The interactions let the first-stage responses of the cell's wage work share and of the market's

[66] Note that identification of lower-nest elasticities also requires that each partition of the wage work sector has within-sector cross-firm variation in tariff reductions, which precludes partitioning the wage work sector by industry.

employment share differ with the cell's characteristics. The exclusion restriction is that the shock and its interactions with baseline characteristics are orthogonal to changes in workers' taste shifters for wage work and for the market.

F.3.4 Reduced-form effects of trade liberalization on market structure

The two derivatives in the policy effect formula come from long-difference regressions at the microregion level,

$$\Delta\widetilde{HHI}_{\bar{g}m} = \beta_{HHI}\,\Delta ICE^*_m + \delta_r + u_m, \qquad \Delta s_m = \beta_{s_m}\,\Delta ICE^*_m + \delta_r + v_m, \tag{F.22}$$

with region fixed effects and 1991 formal employment weights. I compute the expected payroll Herfindahl in RAIS, with each firm's wage bill valued at expected wages, and the wage work share s_m in the census, as the wage bill of formal and informal wage workers over the market's total wage bill, which includes self-employment earnings. I report effects per 10% increase in the shock.

F.3.5 Wage markdowns

I combine the estimated elasticities with the measured market shares to compute the average inverse elasticity of residual labor supply from equation F.10, the wage take-home share $1/(1+\bar{\varepsilon}^{-1}_{\bar{g}m})$, and, with the reduced forms in equation F.22, the effect of trade liberalization on the take-home share. I evaluate all three at the 1991 averages of $\widetilde{HHI}_{\bar{g}m}$ and s_m across microregions, weighted by 1991 formal employment (0.025 and 0.782), and I compute standard errors by the delta method over the two reduced-form coefficients, holding the elasticities fixed.

F.3.6 Specifications behind Table 7

Column (1) of Table 7 reproduces the paper's main estimates, with local labor markets defined as microregion × occupation pairs: Panel A from Appendix Table A.4, column (4), Panel C from Table 2, Panel E from Table 3, and Panel F from Table 4 and Section 7, where the standard error also reflects the uncertainty in the elasticities. Column (2) re-estimates the baseline model with local labor markets defined as microregions, weighted by the microregion's 1991 formal employment, with ΔICE^*_m as the instrument for $1/\theta$. Column (3) reports the extended model with local labor markets defined as microregions, estimated as follows.

Panels A and B. 1991–2000 long differences of the expected payroll Herfindahl and of the wage work share of market payroll on ΔICE^*_m (equation F.22), per 10% increase in the shock, weighted by the microregion's 1991 formal employment, with region fixed effects and standard errors clustered by microregion.

Panel C. 2SLS estimates of $1/\eta$ from 1991–2000 long differences of firm × microregion wage premia on employment (equation F.13), instrumented with the firm-level tariff reduction, with microregion fixed effects and standard errors clustered by firm; first-stage F statistic of 207.1. "Firm Dec. premium" is the firm × market fixed effect on workers' December monthly earnings, estimated conditional on demographics (Section 5).

Panel D. 2SLS estimate of Ω from equation F.17, over 1991–2000, with the instrument of equation F.19 interacted with gender, education, and age group indicators (six instruments remain after the collinear interactions are dropped); first-stage F statistic of 13.5, Hansen J p-value of 0.034. The standard error of $1/\tilde{\eta} = 1/\eta + \Omega$ combines those of its two components as $\sqrt{\mathrm{SE}(1/\eta)^2 + \mathrm{SE}(\Omega)^2}$, treating the two estimates as independent. The unemployment insurance row values four of the twelve months after separation at the formal wage, so that the wage gap is two thirds of its unadjusted value.

Panel E. Joint 2SLS estimates of $1/\tilde{\rho}$ and $1/\tilde{\theta}$ from equation F.21, over microregion × demographic cells with positive 1991 formal employment (7,425 cells), instrumented with ΔICE_m^* and its interactions with age group, gender, education, and the cell's 1991 formal employment share, with region fixed effects, weighted by the cell's 1991 formal employment, with standard errors clustered by microregion; first-stage F statistics of 228.5 for the sector share and 20.3 for the market share, Kleibergen-Paap Wald F of 10.9, Hansen J p-value of 0.0002. "Residual log earnings" are census log monthly earnings residualized on demographics × year at the individual level and averaged to the cell.

Panel F. Equations F.10 and 25 evaluated at the 1991 averages of the expected payroll Herfindahl and the wage work share across microregions, weighted by 1991 formal employment; the standard error is computed by the delta method over the two reduced-form coefficients of Panels A and B, holding the elasticities fixed.

F.4 Symmetric treatment of informal wage work and self-employment

The extended model treats the two informal margins asymmetrically. This subsection considers the alternative in which workers make no distinction between them: each market has two sectors that workers choose between, formal employment, populated by formal firms, and informal employment, which pools informal wage work and self-employment. Workers value formal jobs at posted wages, so the within-market cross-firm elasticity is the baseline η and there is no bias term Ω, and the average inverse elasticity of residual labor supply takes the form of equation F.10, with $\hat{\rho}$ the elasticity of substitution between formal and informal employment within markets, $\hat{\theta}$ the cross-market elasticity, HHI the payroll Herfindahl of formal firms, and the formal share of the market's wage bill s_m^f in

place of s_m. Estimation follows Appendix F.3 with three changes. First, the trade shock is

$$\Delta ICE_m^{\dagger} \equiv s_{m,1991}^{f} \, \Delta ICE_m^{f} + \left(1 - s_{m,1991}^{f}\right) RTR_m^{*}, \tag{F.23}$$

where $\Delta ICE_m^{f} = -\sum_z s_{zm,1991} \Delta \ln(1 + \tau_{i(z)})$ sums over all of the market's formal firms, weighted by their (linear) 1991 payroll shares, with a zero tariff change for non-tradable firms and firms without a tariff match, and RTR_m^{*} is Dix-Carneiro and Kovak (2017)'s exposure measure, RTR, with industry weights computed from the 1991 census employment of informal wage workers and the self-employed only; $\Delta ICE^{\dagger}$ averages 5.6% across microregions. Second, equation F.21 uses the residual earnings of the cell's formal workers and the formal share of the cell's employment in place of the wage work objects, is estimated over the cells for which the cell-level $\Delta ICE^{\dagger}$ is defined (cells with 1991 formal employment in RAIS and tradable-industry informal or self-employment in the 1991 census), weighted by the cell's 1991 population, and instrumented with the cell-level counterpart of $\Delta ICE^{\dagger}$ and its interactions with age group, gender, education, and the cell's 1991 formal employment share. Third, the outcomes of equation F.22 are the payroll Herfindahl and the formal share of the market's wage bill.

Table F.1 reports the results next to those of the extended model. A 10% increase in the trade shock raises the payroll Herfindahl by 0.016 points (SE of 0.004) and lowers the formal share of the market's wage bill by 0.120 (SE of 0.016); the cross-market inverse elasticity is 3.526 (SE of 1.790); the implied take-home share is 53.3 cents on the dollar, and a 10% increase in the trade shock lowers it by 0.37 cents. The inverse elasticity of substitution between formal and informal employment is 0.023 (SE of 0.060): taken at face value, formal and informal employment are close to perfect substitutes. This estimate is biased toward zero if the trade shock moved workers into informal employment within firms (Ulyssea, 2018; Derenoncourt et al., 2025), since the resulting changes in formal and informal employment trace firms' labor demand rather than workers' labor supply. Figures F.3 and F.4 show the pattern: across microregions, larger values of RTR are associated with lower employment shares and lower earnings in wage work, the same-sign pattern of a labor demand shock that traces out labor supply, but with higher employment shares and lower earnings in self-employment and in informal wage work. Absent data that separate the intensive and extensive margins of informality in this period, I cannot estimate the elasticity of substitution between formal and informal wage work from these moments.

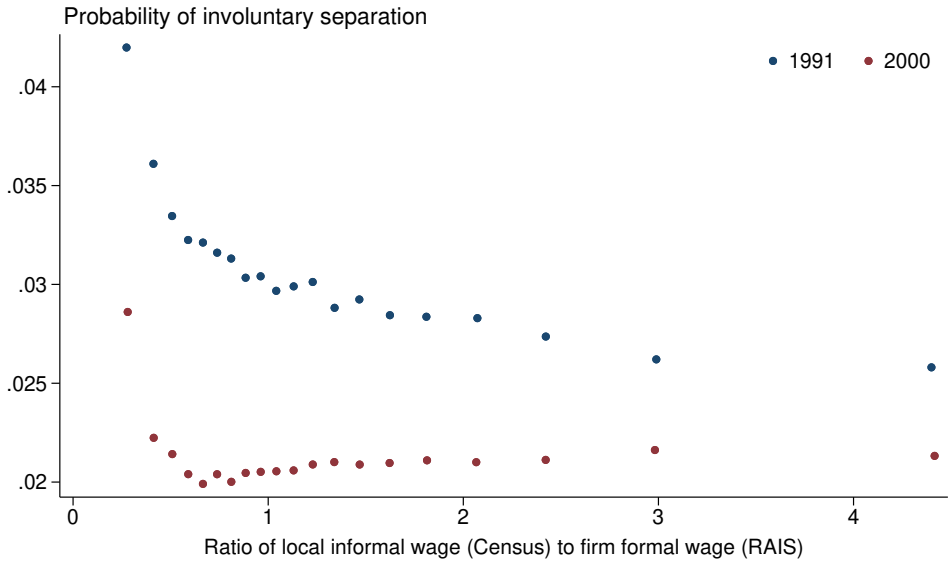
Table F.1: Extended model and symmetric treatment of informal wage work and self-employment

	Extended model (main)	Assuming no preference distinction between informal wage work and self-employment
	(1)	(2)
Panel A: Effect of import competition on formal labor market concentration		
Δ Concentration, per 10% increase in the trade shock	0.028	0.016
	(0.007)	(0.004)
Trade shock	ΔICE^{*}	$\Delta ICE^{\dagger}$
Concentration measure	*Expected* Payroll Herf.	Payroll Herf.
Level of analysis	Mmc	Mmc
Panel B: Effect of import competition on the formal / wage work share of market payroll		
Δ Sector share, per 10% increase in the trade shock	-0.032	-0.120
	(0.005)	(0.016)
Share	Wage work	Formal employment
Trade shock	ΔICE^{*}	$\Delta ICE^{\dagger}$
Level of analysis	Mmc	Mmc
Panel C: Within-market cross-firm inverse elasticity of substitution		
$1/\eta$ (response to formal wages)	0.870	0.870
	(0.062)	(0.062)
Trade shock	Firm-level tariff reduction	
Level of analysis	Firm × mmc	Firm × mmc
Wage measure	Firm Dec. premium (all workers), pair pre-trend controls	
Panel D: Bias Ω in 2SLS estimates of $1/\eta$ due to risk of separation from formal job		
Ω (response to expected formal–informal wage gaps)	0.0091	–
	(0.0025)	
$1/\tilde{\eta} = 1/\eta + \Omega$	0.879	–
	(0.062)	
Trade shock	Firm tariff × ΔICE^{*}	–
Level of analysis	Firm × mmc × demog. cells	–
Fixed effects	Firm, mmc, demographic cell	–
Panel E: Cross-sector and cross-market inverse elasticities of substitution		
Within-market cross-sector inverse elasticity	0.684	0.023
	(0.234)	(0.060)
	$1/\tilde{\rho}$ (wage work vs. self-emp.)	$1/\hat{\rho}$ (formal vs. informal)
Cross-market inverse elasticity	1.187	3.526
	(0.481)	(1.790)
	$1/\tilde{\theta}$	$1/\hat{\theta}$
Trade shock	ΔICE^{*}	Cell $\Delta ICE^{\dagger}$
Fixed effects	Region	Region
Level of analysis	Mmc × demog. cells	Mmc × demog. cells
Wage measure	Residual log earnings, wage workers	Residual log earnings, formal workers
Panel F: Effect of import competition on the average wage take-home share		
Δ average inverse elasticity, per 10% increase in the trade shock	0.0053	0.0132
	(0.0014)	(0.0042)
Implied change in take-home share (pp)	-0.15	-0.37
Baseline (1991) wage take-home share	0.531	0.533
Trade shock	ΔICE^{*}	$\Delta ICE^{\dagger}$

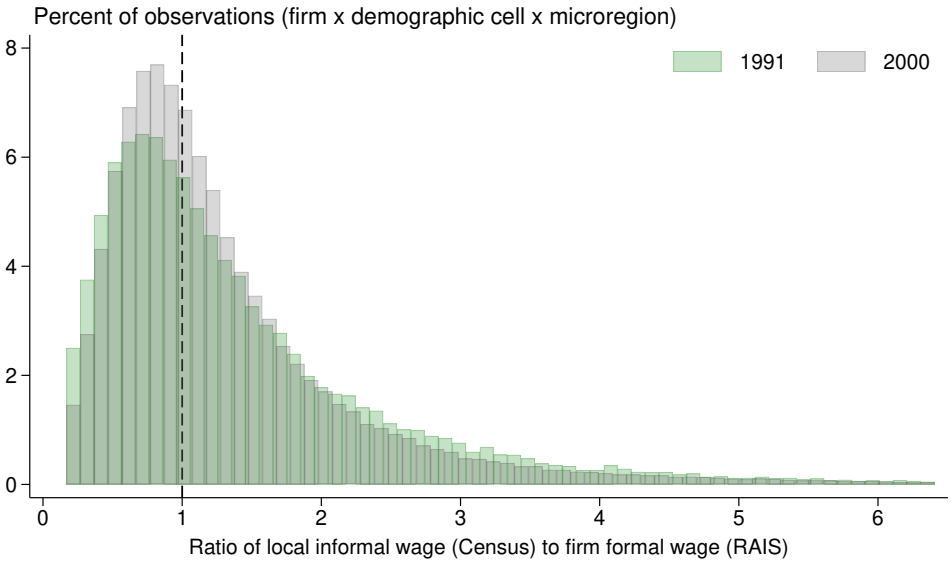
Notes: Standard errors in parentheses. Column (1) repeats column (3) of Table 7. Column (2) reports the model of Appendix F.4, in which informal wage work and self-employment form a single informal sector that workers choose as an alternative to formal employment at posted wages. Specifications are as in Table 7, with $\Delta ICE^{\dagger}$ as the trade shock, the payroll Herfindahl as the concentration measure, the formal share of the market's wage bill as the sector share, and the residual log earnings of formal workers as the wage measure in Panel E. Panel E, column (2), differs in three further respects (Appendix F.4): it covers the cells for which the cell-level $\Delta ICE^{\dagger}$ is defined (6,291 cells in 469 microregions), is weighted by the cell's 1991 population, and is instrumented with the cell-level $\Delta ICE^{\dagger}$ and its interactions. Panels A and B, column (2), include 469 microregions: those of column (1) except the microregions dropped by Dix-Carneiro and Kovak (2017). First-stage F statistics in Panel E, column (2), are 60.7 and 11.2.

Figure F.1: Probability of involuntary separation from formal firm and informal-formal wage gaps

Panel A: Probability of involuntary separation and wage gap



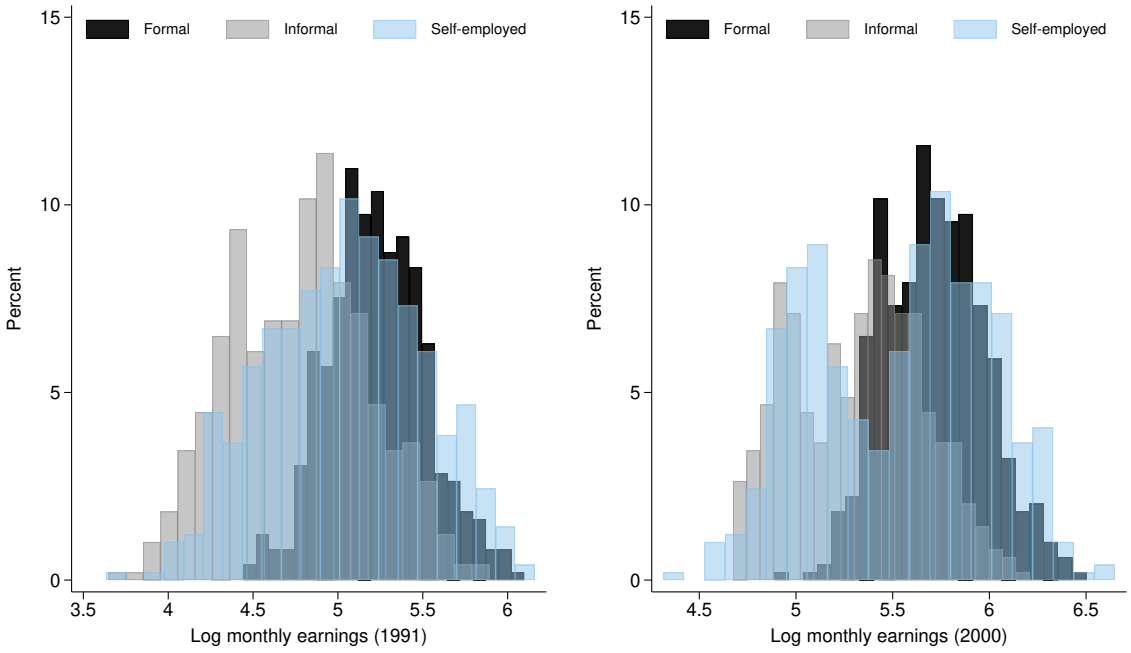
Panel B: Wage gap distribution



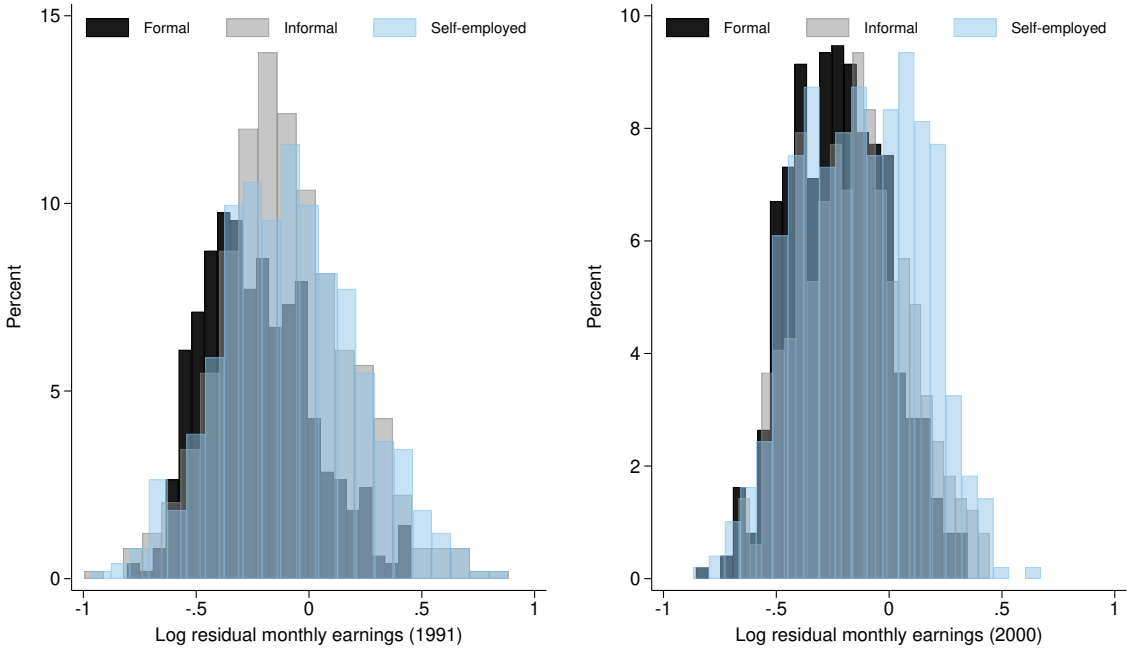
Notes: This figure shows the relationship between probabilities of involuntary separation from a formal sector firm in RAIS and the gap between local informal sector monthly earnings and the firm's formal monthly earnings, separately computed for each demographic cell. The sample includes all formal sector firms in RAIS in 1991 and 2000, merged at the microregion X demographic cell X year level to the corresponding census data. Each observation is a firm X microregion X demographic cell (gender X education group X age group dummies). Monthly earnings gaps are calculated by first converting all reported earnings to 2000 constant Reais. The probability of involuntary separation is calculated using separations between January 1 and December 31 of each year. Firm total formal employment is measured as of December 31 of each year.

Figure F.2: Formal, informal, and self-employment earnings

Panel A: Raw earnings



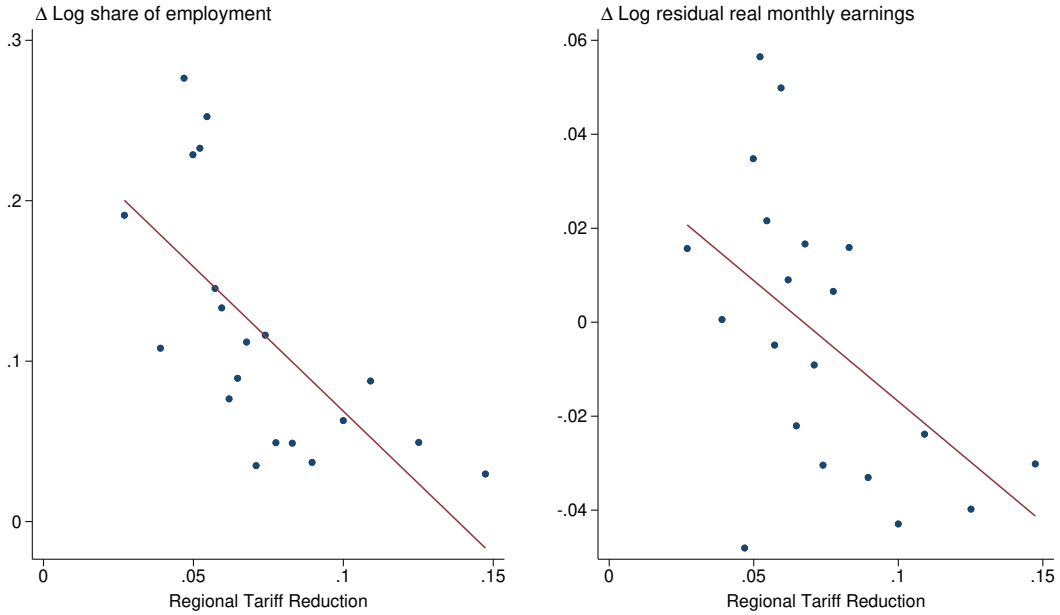
Panel B: Residual earnings



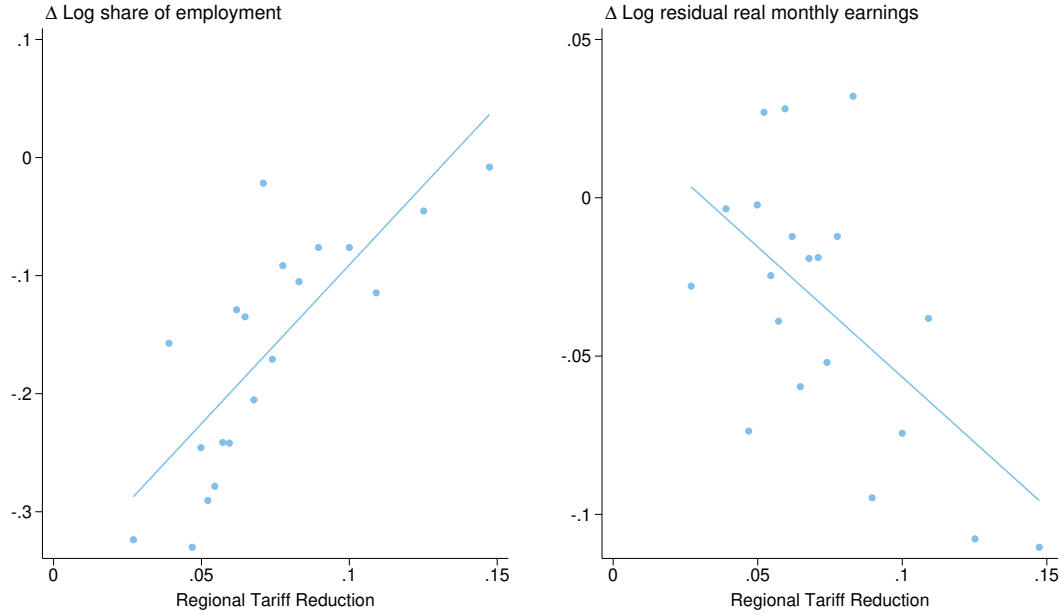
Notes: This figure plots the distribution of raw and residual formal real monthly earnings, informal real monthly earnings, and self-employment real monthly earnings, across 486 Brazilian microregions. Residual real monthly earnings are conditional on a fully saturated vector of dummies for gender, age groups, and education groups. Real earnings are expressed in 2000 reais.

Figure F.3: Effect of Dix-Carneiro and Kovak (2017)'s RTR on wage work and self-employment

Panel A: Wage work



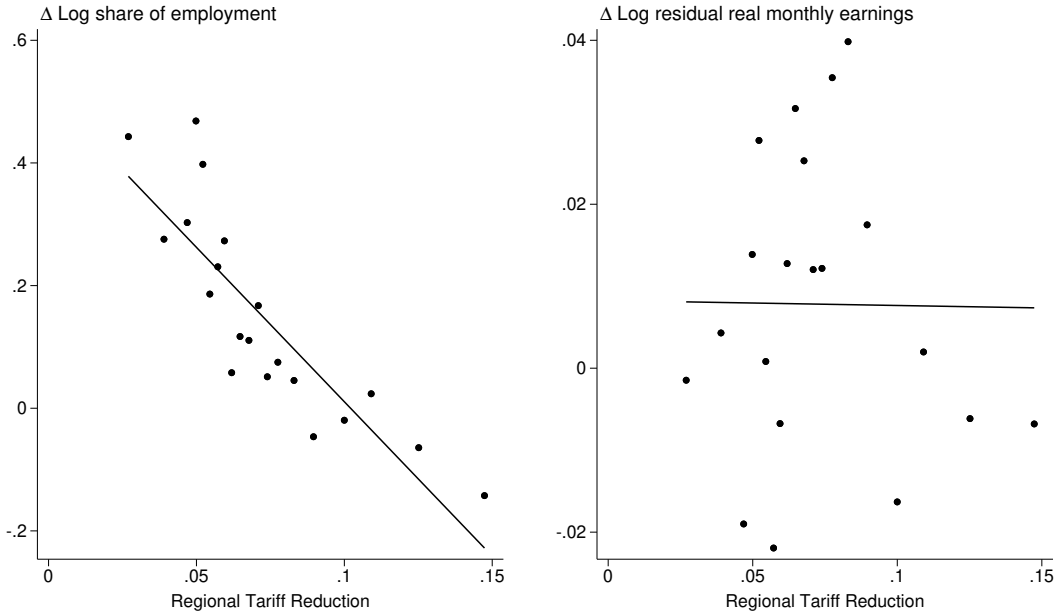
Panel B: Self-employment



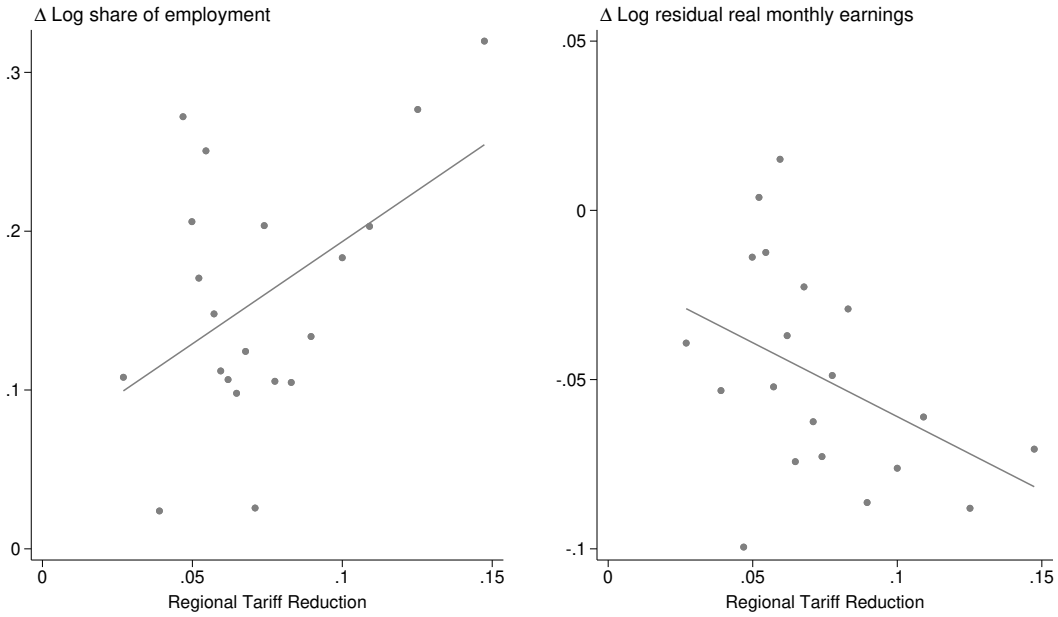
Notes: This figure shows the correlation between Dix-Carneiro and Kovak (2017)'s exposure measure, RTR (regional tariff reductions, i.e., microregion-level exposure to 1990–1994 import tariff reductions), changes in employment shares (left panel), and changes in log residual real monthly earnings (right panel) between the 1991 and 2000 census, for individuals employed in wage work (formal or informal) (Panel A) versus in self-employment (Panel B) across 486 Brazilian microregions. Log residual real monthly earnings are conditional on flexible controls for gender, age, and education. Real earnings are computed from nominal earnings by expressing all values in 2000 reais.

Figure F.4: Effect of import competition on formal versus informal wage work

Panel A: Formal Wage Work



Panel B: Informal Wage Work



Notes: Each panel shows the correlation between Dix-Carneiro and Kovak (2017)'s exposure measure, RTR (regional tariff reductions, i.e., microregion-level exposure to 1990–1994 import tariff reductions), changes in employment shares (left) and log residual real monthly earnings (right) between 1991 and 2000 across 486 Brazilian microregions. Log residual real monthly earnings are conditional on a fully saturated vector of dummies for gender, age groups, and education groups. Real earnings are expressed in 2000 reais.