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Labor Reallocation in Response to Trade Reform*

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Abstract

Tracking individual workers across sectors and firms after Brazil's trade liberalization in the 1990s shows that tariff cuts and additional imports trigger worker displacements, but that neither comparative-advantage sectors nor exporters absorb trade-displaced workers for years. To the contrary, there are more displacements and fewer accessions in comparative-advantage sectors and at exporters, and trade liberalization increases transitions to informal work and self-employment. Labor productivity at exporters increases faster than production so that output shifts to more productive firms but labor does not.

Keywords: International trade; factor reallocation; labor demand and turnover; linked employer-employee data

JEL Classification: F14, F16, J23, J63

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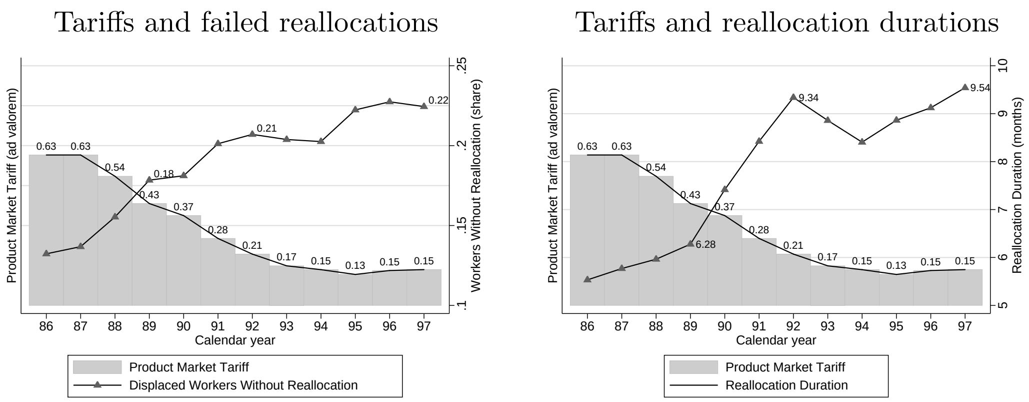
1 Introduction

Economists have long studied the consequences of international trade. Numerous empirical studies investigate the impact of trade on economic outcomes for the country as whole, sectors, firms or plants. Yet research to examine the impact of trade liberalization on workers' individual employment trajectories is scant.[1] We use economy-wide linked employer-employee data and investigate resource reallocation directly, by following workers across employers and industries before and after major trade reform in Brazil. Brazil's trade liberalization triggers worker displacements particularly from protected industries, as trade theory predicts and welcomes. But neither comparative-advantage industries nor exporters absorb trade-displaced workers for years. In fact, comparative-advantage industries and exporters displace significantly more workers and hire fewer workers than the average employer, and resource reallocation remains incomplete for years.

Prior evidence for Brazil and other economies shows that product-market shares are reallocated to more efficient producers, while employers exhibit productivity increases (Pavcnik 2002, Schor 2004, Muendler 2004) and technology upgrading (Verhoogen 2007, Bustos 2005) in response to trade opportunities. Shifts in product-market shares to more advanced firms and exporters are sometimes interpreted as evidence for successful resource reallocation. Brazil's experience shows otherwise. Labor is flowing away from comparative-advantage sectors and away from exporters because their labor productivity increases faster than their production so that output shifts to more productive firms while labor does not.

This labor-market evidence for Brazil also offers a novel explanation why pro-competitive reforms can be associated with strong efficiency gains at the employer level but not in the aggregate. Figure 1 illustrates economic changes in Brazil during the 1990s. Brazil's tariff cuts substantially reduce trade barriers in the early 1990s. At the same time, the share of displaced workers with no reallocation for four years rises from below 18 percent before 1990 to 21 percent by 1993 and the duration of successful reallocations lasts nine months by 1993, up from six months and less before 1990. Conservatively measured, the foregone wage bill from the increase in reallocation durations and failures after 1990 amounts to between one and three percent of GDP. The increase in joblessness is not solely due to trade integration. But regression analysis at the worker level, and a series of robustness checks to rule out alternative explanations, document that trade

[1] Existing worker-level studies compare displaced workers across industries for periods with minor changes to trade and find that employment histories vary little by the displacing sector's trade exposure (Kruse 1988, Hungerford 1995, Kletzer 2001).



Sources: *RAIS* 1986-2001 (1% random sample), male workers nationwide, 25 to 64 years old, displaced from a formal-sector job; not rehired into a formal-sector job within 48 months (*left graph*) or rehired into a formal-sector job within 48 months (*right graph*). Product tariffs from Kume, Piani and Souza (2000), employment weighted at *Nível 50* sector level.

Figure 1: **Tariffs and national labor market performance**

variables predict a large part of the fluctuation in displaced labor.

In the presence of frequent worker transitions between formal employment and other types of work, it takes linked employer-employee data to measure reallocation directions and durations after trade reform. We combine data from numerous sources into a comprehensive data set. From Brazil's labor ministry we gather administrative data with detailed demographic information on every formal-sector worker and the identified employer. At the employer-level, we match information on ownership, labor productivity, and export status. At the sector level, we obtain measures of comparative advantage, tariffs and import penetration. The time dimension of up to sixteen years (between 1986 and 2001) allows us to measure idle resources and to condition on unobserved worker heterogeneity. To assess transitions to informality, self-employment and unemployment directly, we draw on household survey data. We control for concomitant economic changes during the sample period, including macroeconomic stabilization, foreign direct investments, privatization, service-job outsourcing, and a reform of labor-market regulations prior to trade liberalization. We construct instrumental variables for export demand, using sectoral imports from other source-countries than Brazil in foreign destinations and foreign price components of the sectoral real exchange rate. The instruments are strong predictors of trade policies and export behavior, and rule out that simultaneity between labor turnover and trade adversely affects our estimates.

The empirical literature on trade and resource reallocation has taken three main approaches. First, industry-level studies use measures of job creation, destruction, and churning (excess turnover beyond net change), as well as informality. Haltiwanger, Kugler, Kugler, Micco and Pagés (2004) show for a panel of six Latin American countries, for instance, that tariff reductions are associated with heightened within-sector churning and net employment reductions at the sector level.[2] Beyond those studies, linked employer-employee data in our paper document the direction of factor flows between types of employers, and identify the incidence of idle resources in the process. In contrast to the United States, where industries with faster productivity growth exhibit higher net employment growth (Davis et al. 1996), more productive employers reduce employment in Brazil during the 1990s. Using sector data, Goldberg and Pavcnik (2003) report no statistically significant relation between informal work and trade in Brazil, whereas household survey data in our paper document that tariff reductions predict significantly more transitions to informal work.

Second, employer-level studies show that trade reforms are associated with product-market reallocation towards more efficient producers (for a survey see Tybout 2003). But employer-level studies typically report no detectable relationship between trade and employment.[3] In our data, trade variables are not statistically significant predictors of employment changes at the employer-level either (Muendler 2007). But worker-level regressions in this paper, on the same data, uncover that additional imports trigger significantly more worker displacements, while there are lasting worker flows away from productive high-output employers. This suggests that unobserved workforce heterogeneity hampers regressions at more aggregate levels, even the employer level, and calls for the use of worker panel data.

[2]Using measures of net employment change, Wacziarg and Wallack (2004) detect no statistically significant labor reallocation in a cross-country cross-sector study of trade-liberalization episodes. Other examples of industry-level studies include Davis, Haltiwanger and Schuh (1996) for the United States, Roberts (1996) for developing countries, and Ribeiro, Corseuil, Santos, Furtado, Amorim, Servo and Souza (2004) for Brazil.

[3]Roberts (1996) reports no clear effect of time-varying trade exposure on employment changes at plants in Chile and Colombia when sector characteristics are taken into account. Using Chilean plant data, Levinsohn (1999, p. 342) concludes that, "try as one might, it is difficult to find any differential employment response" to trade liberalization. Neither do Davis et al. (1996) find a clear effect of trade on gross job flows using U.S. data. An exception is Biscourp and Kramarz (2007) who show that French firm-level trade data exhibit a significant association of job destruction with firm-level imports.

Third, a worker-level literature studies the experience of displaced workers across sectors and worker groups. Kruse (1988) and Kletzer (2001) compare displaced workers between U.S. industries and find that employment histories are largely explained by differences in workforce characteristics across sectors and vary little by a sector's trade exposure.[4] Time variation in our data, by contrast, identifies a salient impact of Brazil's trade opening on labor turnover. Beyond displaced-worker survey data, our linked employer-employee records allow us to quantify directions of worker flows across employers for many years and show that the economic burden of trade-induced joblessness is substantial.

The paper has seven more sections. Section 2 discusses the data (with some details relegated to the Appendix). Section 3 reports descriptive evidence on trade and labor reallocation in Brazil. Section 4 presents estimates of work status transitions from a household perspective. Section 5 analyzes worker separations and accessions to identify sector and firm predictors that explain reallocation delays and failures. Section 6 subjects these predictions to numerous robustness checks. Section 7 discusses the implications for trade theory. Section 8 concludes.

2 Data

We track Brazil's labor reallocation with two main data sets. First, we construct linked employer-employee data on the basis of Brazil's comprehensive labor force records *RAIS* for the 16-year time span from 1986 through 2001. The *RAIS* data include all formally employed workers, identify their employers and industries, and track the workers over time so that worker-fixed effects are estimable and national formal-sector migration is covered.[5] To *RAIS*, we match information on the employer's export status from national customs records and sector-level information from various sources. By design, however, workers with no current formal-sector employment are not in *RAIS*. So, for our second data source, we use the metropolitan household survey *PME*. *PME* provides direct information on household members with or without formal-sector employment and covers one work status transition at the annual horizon for every household member.

[4]Similarly, Hungerford (1995) finds that short-term trade shocks play a minor role for separation rates in the United States.

[5]Formal-sector migration is substantial, especially across metropolitan areas. Among the prime-age male workers in *RAIS* with a metropolitan job in 1990, for instance, 15 percent have a formal job outside the 1990 city of employment by 1991 and 25 percent by 1993. Similarly, among the metropolitan workers in 1994, 17 percent have a formal job in another city by 1995 and 27 percent by 1997.

Linked employer-employee data. *RAIS* (*Relação Anual de Informações Sociais*) provides comprehensive annual information on workers formally employed in any sector (including the public sector). Our full data include 71.1 million workers with 556.3 million job spells at 5.52 million plants in 3.75 million firms between 1986 and 2001.[6] Every job observation is identified by the worker ID (PIS), the plant ID (of which the firm ID is a systematic part), the month of accession, and the month of separation. Relevant worker information covers educational attainment in nine categories, tenure at the plant, age, and gender; job information includes the monthly average wage and an occupation classification comparable to the ISCO-88 four-digit level; spell classifications report reasons for separations and accessions as well as contractual arrangements; plant information includes sector, municipality, and public-private ownership categories (see the Appendix for details). We infer a plant's workforce composition by aggregating *RAIS* to the plant level.

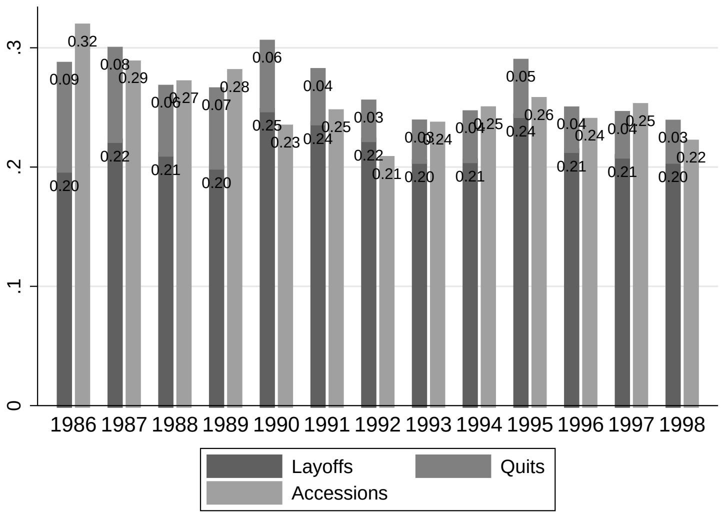
To construct the worker sample, we take the list of all proper worker IDs (11-digit PIS) that ever appear in *RAIS*, draw a one-percent random sample of the IDs, and then track the selected workers through their formal jobs. Industry information in this paper is based on the subsector IBGE classification (roughly comparable to the NAICS three-digit level), which is available by plant over the full period.[7]

We report results for prime-age males, 25 to 64 years old, in order to focus on workers after their first labor-force entry and to be little affected by labor supply changes.[8] For most separation statistics, we remove multiple jobs and

[6] *RAIS* provides information to the Brazilian labor ministry MTE primarily for a federal wage supplement program (*Abono Salarial*), by which every worker with formal employment during the calendar year receives the equivalent of a monthly minimum wage. *RAIS* records are then shared across government agencies. An employer's failure to report complete workforce information can result in fines proportional to the workforce size; but fines are seldom issued. In practice, workers and employers have strong incentives to ascertain complete *RAIS* records because payment of the annual public wage supplement is exclusively based on *RAIS*. The ministry of labor estimates that currently 97 percent of all formally employed workers in Brazil are covered in *RAIS*, and that coverage exceeded 90 percent throughout the 1990s.

[7] We also draw a five-percent sample of all worker IDs that ever appear in a metropolitan area for direct comparisons to *PME*, and follow the workers nationwide. In addition, we repeat the calculation of statistics and estimation with the CNAE sector classification (roughly comparable to the NAICS four-digit level), which is available since 1995, by using a sample of 1995-survivor plants. The additional figures, tables and estimates are available online at *wwww.econ.ucsd.edu/muendler/research*. Results exhibit little sensitivity to alternative regional samples or sector classifications.

[8] Results are similar for samples that include both genders and all age groups.



Source: *RAIS* 1986-98 (1% random sample), male workers nationwide, 25 to 64 years old, with employment in subsector IBGE manufacturing (highest paying job if many). Separation and accession rates exclude transfers, deaths, and retirements and are relative to totals of first and last observed employments in a given year.

Figure 2: **Separations and accessions in manufacturing**

only retain a worker's highest paying job at a given moment. When we infer separations and accessions, we exclude transfers across plants within the same firm, as well as retirements and reported deaths on the job. For much of the analysis, an accession is defined as a worker's hiring into the first employment in the calendar year; reference observations to accessions are jobs with no reported accession at the hiring employer in the year. Conversely, we define a separation as a worker's quit or layoff from the last employment in the calendar year; reference observations for separations are jobs with no reported separation from the firing employer in the year.

Displaced workers who await reallocation are not directly observable in *RAIS*. However, we observe changes to the pool of workers *to be reallocated* at two margins: separations from formal jobs add to the pool of workers to be reallocated, and re-accessions into formal jobs shrink the pool of workers to be reallocated. Figure 2 documents changes in gross separation and accession rates. In manufacturing, annual accession rates exceed separation rates before 1990 (except for 1987). During the years of trade reform 1990-91, separation rates rise. At the same time, hiring rates exhibit a marked drop from .28 in 1989 to .23 in 1990, and subsequently remain at or below .25 throughout 1997. Except for 1994 and 1997, separation rates are below accession rates after 1991. Our analysis focuses on these two margins of formal-sector labor force adjustment. Among the separations, quits are infrequent compared to layoffs (Figure 2). We mostly consider separations as a single category.

Table 1 shows in the upper two panels, first for the universe of workers and second for prime-age male workers, that the changes to gross separation and accession rates in manufacturing result in a net labor-force shift out of manufacturing over the course of the 1990s. While the manufacturing sector employs 24 (27) percent of Brazil's national formal workforce in 1990, the share is only 19 (21) percent by 1998. Agriculture, commerce and services employ larger shares in 1998 than in 1990. Construction, utilities, and public administration are shown as other sectors in Table 1 and exhibit a reduction in relative importance between 1990 and 1998. Our prime-age male sample represents a population of around 10.8 million workers in Brazil in 1990 and 11.6 million in 1998. Allocation shares and changes for prime-age male workers are similar to the universe of workers. The third panel of Table 1 presents employment shares for the metropolitan areas of Brazil and shows that metropolitan labor markets exhibit employment shares and undergo changes in relative employment similar to the economy as a whole. More than half of Brazil's male formal-sector workforce is employed in metropolitan areas.

Metropolitan household data. The Brazilian monthly employment survey *PME* (*Pesquisa Mensal de Emprego*) provides details on work status. *PME* data derive from a random sample of households in six metropolitan areas (São Paulo, Rio de Janeiro, Belo Horizonte, Porto Alegre, Salvador, Recife). The data are collected by Brazil's statistical bureau IBGE from a rotating panel similar to the U.S. *PSID*. *PME* follows households for 16 months, with an eight-month interval after the fourth interview.[9] Changes to the sample design adversely affect worker panels starting in odd years. So, we use only individuals whose first survey occurs in 1986, 1988, 1990, 1992, 1994, 1996 or 1998. As with *RAIS*, we restrict our sample to prime-age male workers. We trace work-status changes at the annual horizon between the fourth and the eighth interview for each household member and control for the individual's work status during the three months prior to the

[9]Individuals within households are surveyed for a total of eight interviews over a 16-months period. Denoting the initial month with m, interviews are at m, $m+1$, $m+2$, $m+3$, $m+12$, $m+13$, $m+14$, and $m+15$.

Table 1: EMPLOYMENT BY SECTOR AND FORMALITY STATUS

	Agric.	Manuf.	Comm.	Services	Other	*Total*[a]
	(1)	(2)	(3)	(4)	(5)	(6)
			Allocation nationwide (*RAIS* universe)			
1990	.016	.240	.128	.278	.338	23,174
1998	.041	.187	.154	.320	.299	24,492
			Allocation nationwide (*RAIS* prime-age males)			
1990	.019	.273	.111	.284	.314	10,763
1998	.057	.214	.134	.308	.286	11,640
			Allocation in metropolitan areas (*RAIS* prime-age males)			
1990	.007	.277	.104	.309	.302	5,965
1998	.017	.203	.125	.369	.285	6,057
			Informality in metropolitan areas (*PME* prime-age males)			
1990	.159	.063	.109	.117	.298	
1998	.232	.120	.154	.169	.341	

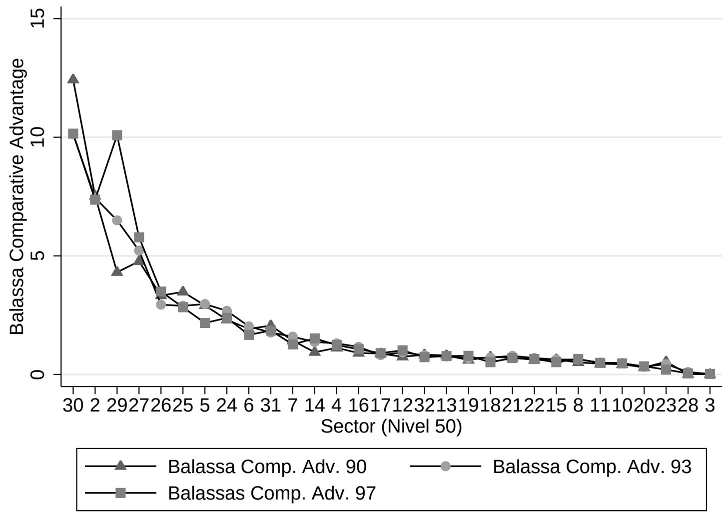
[a]Total employment (thousands of workers), samples scaled to population equivalents.

Sources: *RAIS* 1990 and 1998, employed on December 31; if indicated, male workers nationwide (1% random sample) and in metropolitan areas (5% random sample), 25 to 64 years old. *PME* 1990 and 1998, male workers 25 to 64 years old, and employed at September interview. Manufacturing includes mining.

fourth interview.

PME distinguishes formal employment (with a labor ID card, *carteira*) and informal employment (without ID card). The ID card entitles workers to benefits mostly borne by the employer. In the survey, individuals without employment are considered unemployed if they report active search for work during the week prior to the interview, and are considered out of the workforce otherwise. Household members who work for their own account but do not employ others are considered self-employed. We exclude individuals who become employers.

The lower panel of Table 1 shows that there is a marked increase in informal work status over the 1990s across all sectors. By far the strongest relative increase in informality occurs in manufacturing, where the share of informal workers almost doubles from above 6 to 12 percent. Non-manufacturing industries exhibit an average increase in informality of around fifty percent.



Source: UN Comtrade 1986-98. Sectors at *Nível 50* ranked by sector-fixed Balassa comparative advantage (for sector codes see Table 15 in the Appendix). Estimates of sector-fixed Balassa comparative advantage from regression of Balassa advantage on sector indicators, year indicators, output tariffs and input tariffs between 1990 and 1998.

Figure 3: **Balassa Comparative Advantage**

Sector data. We combine sector-level variables from several sources with *RAIS* and *PME*. We calculate Balassa (1965) comparative-advantage measures for Brazil from UN Comtrade trade data for 1986-98. Sector i's Balassa advantage in year t is

$$BADV_{i,t} \equiv \frac{X_{i,t}^{\text{Brazil}} / \sum_k X_{k,t}^{\text{Brazil}}}{X_{i,t}^{\text{World}} / \sum_k X_{k,t}^{\text{World}}},$$

where $X_{i,t}$ are exports. Note that this index measures revealed comparative advantage from international comparisons of exports data, and is blind to possible sources of advantage. Any explanation of comparative advantage is consistent with this measure.

Brazil's comparative advantage pattern is remarkably stable over the sample period. Figure 3 ranks manufacturing industries by their sector-fixed component. The sector-fixed component is from a linear regression of $BADV$ on sector indicators, year indicators, and product and input tariffs for the years 1990-1998. $BADV$ is not statistically significantly related to tariffs in regressions; and year indicators are neither individually nor jointly different from zero at common significance levels.[10] Figure 3 illustrates the regression results. With the exception of processed sugar (sector 29), Brazil's comparative advantage changes hardly at all. Removal of the sugar-processing sector from our regressions shows that results are not sensitive to its inclusion. Overall, the sector ranking by comparative advantage remains largely unaltered over time.

Our main instrumental variables for export status are imports into Brazil's export destinations from countries other than Brazil, weighted with Brazil's sectoral export volumes in the base year 1990. We use *WTF* (NBER) data on bilateral trade 1990-98 to construct the instruments by subsector IBGE and seven world regions.[11] Additional instruments are components of the sectoral real exchange rate: the U.S. dollar exchange rate and sector price levels in the United States and the EU. We denote exchange rates as foreign divided by domestic prices so that a high exchange rate means a depreciated level. Tariffs are measured as *ad valorem* shares of the import value and market-penetration rates as the share of imports per domestic absorption. We relegate further details on the instruments and sector variables to the Appendix.

Firm data. We combine the linked employer-employee data from *RAIS* with additional firm-level data.[12] Annual customs office records on exports are available to us from *SECEX* (*Secretaria de Comércio Exterior*) for 1990 through 1998. We set the indicator variable for a firm's exporting status to one iff *SECEX* records show exports of any product from the firm in a given year.[13] We link the export-status indicator to *RAIS* at the firm level. For select robustness checks, we also link firm-level labor productivity from the manufacturing survey *PIA* to *RAIS*.

[10] Results are online at *www.econ.ucsd.edu/muendler/research*.

[11] We calculate sector-specific weights for each foreign destination country in 1990 using *SECEX* exports data for Brazil (based on *RAIS* sector information for the *SECEX* exporters). We then calculate aggregate imports into each foreign country, excepting imports from Brazil, and weight the country aggregates with Brazilian export volumes by sector and destination in 1990. We finally aggregate the sector-weighted country totals to seven world regions and obtain seven foreign import-demand instruments that vary by sector and year. The seven world regions are Asia-Pacific Developing countries (APD), Central and Eastern European countries (CEE), Latin American and Caribbean countries (LAC), North American countries (NAM excluding Mexico), Other Developing countries (ODV), Other Industrialized countries (OIN), and Western European countries (WEU).

[12] There are no employer identifiers in the *PME* household survey.

[13] We do not use sales thresholds to define the export indicator because sales information is only available for a small subsample of (*PIA*) firms.

Table 2: PRODUCTIVITY VARIATION ACROSS FIRMS AND OVER TIME

	TFP and Output shares				Labor Prod. and Employment shares			
	Cross section			Ann. chg.	Cross section			Ann. chg.
	wgtd.	unwgtd.	cov.	raw cov.[a]	wgtd.	unwgtd.	cov.	raw cov.[a]
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
1986	1.018	.924	.095		1.011	1.019	-.008	
1990	1.000	.899	.101	.065	1.000	.997	.003	-.029
1992	1.017	.911	.105	.075	1.015	1.008	.007	-.058
1994	1.013	.918	.096	.067	1.023	1.019	.005	-.043
1998	1.035	.910	.125	.047	1.073	1.043	.030	-.039

[a]Four-year lagged average of raw covariances between annual share and outcome changes.

Source: *PIA* firms 1986-98 (1991 missing); log total factor productivity from Muendler (2004) based on Olley and Pakes (1996) estimation (at *Nível 50*), inferring labor productivity under changing capital stocks and intermediate-input uses. Cross-sectional productivity decomposition as in Olley and Pakes (1996): $y_t = \bar{y}_t + \sum_i \overline{\Delta}\theta_{it}\overline{\Delta}y_{it}$, where y_t is weighted and $\bar{y}_t$ is unweighted mean log productivity and $\overline{\Delta}$ denotes deviations from cross-section means (rebased to unity in 1990). Annual productivity change correlation $\sum_{i\in C}\Delta\theta_{i,t}\Delta y_{i,t}$ (raw covariance) from Haltiwanger (1997) decomposition, where Δ denotes annual change (not rebased).

Details are discussed in the Appendix.

3 Labor Reallocation and Trade

Labor reallocation is the reassignment of workers to jobs across employers and sectors. We turn to descriptive evidence on labor reallocation and its relation to Brazil's trade reform and other economic changes between 1986 and 1998.

Labor and output reallocation. In the presence of firm-level productivity change and exit, labor reallocation is distinct from the reallocation of product market shares. If a firm's labor productivity rises faster than its output, additional output is associated with less employment. Similarly, if firms exit but survivors and entrants raise labor productivity faster than output, output shares are being reallocated while labor reallocation remains incomplete. Product-market reallocations to more productive firms and simultaneous workforce shifts away from more productive firms are thus a theoretical possibility; they are Brazil's reality during the 1990s.

Table 2 decomposes total factor productivity (columns 1-4) and labor productivity (columns 5-8) into the contributions of firm-level productivity and firm-level weights, where the weights are output in the case of total factor productivity and employment in the case of labor productivity. The statistics are based on output and employment at formal-sector manufacturing firms. Following Olley and Pakes (1996), aggregate productivity in the cross section of firms (columns 1 and 5) is split into the unweighted mean productivity level (columns 2 and 6) and the covariance between deviations of the weights and productivities from annual means (columns 3 and 7). The rebased overall TFP gain of 3.5 percent between 1990 and 1998 is modest (column 1).[14] Substantial capital accumulation contributes to the faster increase in log labor productivity by 7.3 percent between 1990 and 1998 (column 5). Alongside, Table 2 reports the raw covariance of year-over-year productivity changes at surviving firms (columns 4 and 8)—a term in the Haltiwanger (1997) decomposition over time.[15]

The decompositions in Table 2 show for the cross section of Brazilian manufacturers that firms with higher total factor productivity (TFP) do command larger output shares (column 3), and that TFP improvements among survivors are associated with gains in output shares (column 4). These facts are well known for Brazil and similar countries, but often confounded with resource allocation. The cross-sectional covariance between labor productivity and employment shares, in fact, is considerably weaker (column 7) than between TFP and output shares (column 3). Most strikingly, firm-level labor productivity advances are associated with reductions in employment shares (column 8).[16] So, firms with increasing productivity expand output shares but reduce employment. Resource reallocation is distinct from output reallocation.

Economic reforms. In 1990, the Brazilian government breaks with the country's decade-old import substitution policy and embarks on drastic trade liberalization. Tentative *ad valorem* tariff reductions during the late 1980s were rendered largely ineffective because of binding non-tariff barriers (Kume, Piani and Souza 2000). By contrast, far-reaching trade reform under the Collor administration in

[14] In Table 2, we divide aggregate log productivity levels by the aggregate 1990 log level. Rebasing to 1986 at the firm level in Muendler (2004) yields a 4.7 percent increase between 1990 and 1998.

[15] Centered covariances exhibit a similar pattern as the raw covariances, with always positive TFP and always negative labor productivity covariations. To facilitate comparisons to other research, we report the raw covariance from the Haltiwanger decomposition.

[16] It is mostly firm exits that raise the covariance between labor productivity and employment in the cross section over time (column 7).

1990 involves both the removal of non-tariff barriers and the adoption of a new tariff structure. Collor abolishes all non-tariff barriers by presidential decree on his first day in office. Implementation of the new tariff structure with lower levels and less cross-sectoral dispersion is mostly complete by 1993. Figure 1 above documents the drop in product tariffs from an average level of 63 percent in 1987 to 15 percent by 1997. The new tariff structure also reduces the cross-sectoral dispersion. While product tariffs range between 21 (metallic products) and 63 percent (apparel and textiles) in 1990, they drop to a range spanning between 9 percent (chemicals) and 34 percent (transport equipment) in 1997. Manufacturing industries receive effective protection in both years. In 1990, product tariffs are around 45 percent above intermediate-input tariffs in value-added terms. By 1997, however, the reduced cross-sector dispersion of tariffs results in a smaller rate of effective protection of about 20 percent on average.

Additional reforms partly coincide with trade liberalization. Privatization efforts for public utilities begin in the early 1990s and accelerate by the mid 1990s, while Brazil simultaneously removes capital-account restrictions. In 1994, drastic anti-inflation measures succeed for the first time in decades. These reforms are accompanied by a surge of foreign direct investment inflows during the mid 1990s and advances in outsourcing of service jobs across domestic employers. The pro-competitive product-market reforms of the 1990s were preceded by a labor-market reform in 1988: Brazil's new constitution introduced a series of changes that reduced the work week and increased overtime premia and workers' benefits—significantly raising labor costs (Paes de Barros and Corseuil 2004). Concomitant reforms not withstanding, its scope and pace make trade liberalization a focal candidate to explain employment shifts out of manufacturing and work status transitions into informality.

Workforce characteristics and trade exposure. Table 3 provides a summary comparison of variables for manufacturing industries in different quintiles of comparative advantage, and between exporters and the average employer. Top comparative-advantage industries (in the highest quintile) show a higher labor turnover than the average sector with both more worker separations and more accessions, whereas exporting firms exhibit below-average turnover with both fewer worker separations and fewer accessions than average. Among the separations, reported quits play a minor role.

The average exporter is active in a sector with a slightly lower than average comparative advantage level. Similarly, there are fewer worker observations

Table 3: *RAIS* Summary Statistics for Manufacturing

	All sectors and firms		5th comp. adv. quintile	Exporter
	Mean	Std.Dev.	Mean	Mean
	(1)	(2)	(3)	(4)
Outcomes				
Indic.: Separation	.282	.450	.314	.260
Quit	.026	.160	.031	.020
Indic.: Accession	.292	.455	.326	.237
Main covariates				
Balassa (1965) Comp. Adv.	1.450	1.047	3.223	1.373
Exporter Status	.495	.500	.439	1.000
Product Market Tariff	.193	.103	.174	.204
Intm. Input Tariff	.146	.077	.105	.154
Import Penetration	.064	.052	.031	.074
Plant-level covariates				
Log Employment	5.148	1.952	5.551	6.210
Log Employment 1998/90	.930		.919	.976
Log Labor Productivit	11.186	.706	11.081	11.233
Log Labor Productivity 1998/90	1.045		1.025	1.047

Source: *RAIS* 1990-98 (1% random estimation sample), male workers nationwide, 25 to 64 years old, with manufacturing job. Statistics based on separation sample, except for accession indicator (146,787 observations in separation, 112,974 in accession sample). Sector information at subsector IBGE level. *PIA* 1986-98 for labor productivity information.

at exporters in a top comparative-advantage sector than at exporters overall. The reason is that there is a larger number of small-scale exporters in industries without comparative advantage.[17] Expectedly for a country with a history of import-substitution industrialization, Brazil's top comparative-advantage industries have lower-than-average tariffs. Comparative-advantage industries also exhibit lower import penetration. Firms in top comparative-advantage industries and exporters have larger workforces than average (85 and 326 workers more, respectively, than the average formal-sector manufacturing plant with 257 workers). Recall that our sample is a random draw of workers from the formal-sector worker universe so that larger plants are more likely to be included. Manufacturing employment drops between 1990 and 1998, and drops faster than average in the highest-quintile advantage sectors.

For labor productivity and several additional characteristics (not reported here

[17]We control for employment in our regressions to capture exports-per-worker effects.

Table 4: Annual Sector Transitions and Failed Re-Accessions

To:	Traded: Comp. adv. quintile[a]					Nontraded	*Failure*	*Total*
	1st	2nd	3rd	4th	5th			
From: (in %)	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Traded: Comp. adv.[a]								
1st quintile	14.6	7.4	3.1	6.2	2.8	35.3	30.7	100.0
2nd quintile	6.5	14.2	3.3	4.6	3.3	35.7	32.5	100.0
3rd quintile	3.2	3.6	14.2	7.1	2.8	34.5	34.5	100.0
4th quintile	2.1	2.1	2.7	26.3	5.5	28.3	33.2	100.0
5th quintile	1.9	2.7	1.7	11.2	19.5	32.5	30.4	100.0
Nontraded	1.3	1.5	1.3	3.3	1.8	57.9	32.9	100.0
Failure	3.0	3.1	3.4	11.3	5.0	74.1	.0	100.0
Total	2.6	2.7	2.7	8.4	4.0	60.6	19.1	100.0

[a]Balassa (1965) comparative advantage, transition year quintile (5th: strongest advantage).

Source: *RAIS* 1986-2001 (1% random sample), male workers nationwide, 25 to 64 years old. UN Comtrade 1986 for Balassa comparative advantage; defined at two-digit sector level (Subsector IBGE). Transition frequencies are job accessions in Brazil within one year after separation, based on last employment of year (highest paying job if many). Failed accessions are separations followed by no formal-sector accessions anywhere in Brazil within a year, excluding workers with prior retirement or death, or age 65 or above on earlier job.

for brevity), there are remarkable mean differences between an exporter and an average firm in a top comparative-advantage sector.[18] The reason is that substantial employer heterogeneity prevails within industries, with diverse exporters and nonexporters shifting mean characteristics. Labor productivity increases between 1990 and 1998. At exporters, labor productivity is higher than average over the whole sample period, but lower than average at firms in comparative-advantage industries. Log labor productivity in 1998 exceeds log labor productivity in 1990 by 4.5 percent in the estimation sample, and by 4.7 percent at manufacturing exporters.

Worker reallocation. Table 4 reports transitions of displaced prime-age male workers from formal-sector jobs to other formal-sector jobs at the annual horizon (columns 1-6) for the period 1986-2001, and the share of displaced workers with no observed formal-sector rehiring within a year (column 7). Retained workers do not enter the statistics. Agricultural, mining and manufacturing plants are

[18]Summary statistics for all regressors are in our working paper (Menezes Filho and Muendler 2007).

grouped into their sector's comparative advantage quintiles at the subsector IBGE level.[19] All other sectors—commerce, services, construction, utilities, and public administration—are considered nontraded for the purposes of Table 4.

The majority of successful worker reallocations within traded-goods sectors is to employers in the same comparative-advantage quintile: transition rates along the diagonal in the five traded-merchandize sectors exceed those off the diagonal (column 1-5). Transitions to sectors with similar comparative advantage occur more frequently than to dissimilar sectors: off-diagonal entries are small, especially for accession sectors whose comparative advantage rank is two or more quintiles away from the separation sector. These facts suggest that traded-goods sectors with different degrees of comparative advantage are little permeable to labor reallocation. Classic trade theory posits, in constrast, that factors are reallocated from traded-goods industries with a comparative disadvantage to traded-goods industries with comparative advantage so that the largest fraction of reallocated workers should move to the high-quintile industries (columns 4 and 5) from every separation sector. Only in the aggregate of all separations (last row), including reallocations that failed at the annual horizon before, is there a higher absorption rate into comparative advantage industries.

The dominant fraction of workers with displacement from a traded-goods industry, about a third, finds employment in nontraded-output sectors (column 6). And almost as many workers with displacements from a traded-goods sector, roughly another third, are not rehired into any formal job within a year (column 7). Three out of four workers who are not reallocated at the annual horizon, but who find reemployment in subsequent years, move to the nontraded sector (second-to-last row) and, among the traded industries, mostly into high-quintile industries. Repetitions of the statistical exercise for various subperiods show that annual reallocation patterns in traded-goods industries are remarkably stable. At the four-year horizon, rates of failed transitions drop by around 7 percent in favor of an increase of reallocations along the traded-sector diagonal, whereas off-diagonal entries and transitions to non-traded sectors remain roughly the same (Table 5). The patterns are broadly consistent with the idea that work status changes out of formality (recorded as failures here), and jobs in nontraded-output sectors, provide a buffer for labor reallocation after trade reform.

Linked employer-employee data allow us to track worker reallocations across

[19] Statistics for a sample of 1995-survivor plants and the CNAE sector classification (roughly comparable to the NAICS four-digit level), which is available since 1995, exhibit no noteworthy difference.

Table 5: FOUR-YEAR SECTOR TRANSITIONS AND FAILED RE-ACCESSIONS

To:	Traded: Comp. adv. quintile[a]					Nontraded	*Failure*	*Total*
	1st	2nd	3rd	4th	5th			
From: (in %)	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Traded: Comp. adv.[a]								
1st quintile	23.7	7.5	3.2	8.8	2.9	30.9	23.0	100.0
2nd quintile	8.5	20.2	3.1	6.4	4.2	33.9	23.7	100.0
3rd quintile	4.0	4.1	17.2	12.8	2.4	31.6	27.9	100.0
4th quintile	3.8	3.7	9.1	25.2	5.2	29.3	23.7	100.0
5th quintile	2.3	3.0	2.3	12.9	23.5	33.4	22.5	100.0
Nontraded	1.7	1.7	1.8	4.5	2.6	57.8	29.9	100.0
Failure	3.7	3.0	5.2	15.0	7.1	66.1	.0	100.0
Total	3.8	3.3	4.3	11.0	5.5	56.2	16.0	100.0

[a]Balassa (1965) comparative advantage, transition year quintile (5th: strongest advantage).

Source: *RAIS* 1986, 1990, 1994 and 1998 (1% random sample), male workers nationwide, 25 to 64 years old. UN Comtrade 1986 for Balassa comparative advantage; defined at two-digit sector level (Subsector IBGE). Transition frequencies refer to employments in Brazil four years after separation, based on last employment of year (highest paying job if many). Failed accessions are separations followed by no formal-sector employment anywhere in Brazil after four years, excluding workers with retirement or death, or age 65 or above in past job.

identified firms. Table 6 shows flows of prime-age male workers between nonexporters and exporters for the two periods 1990-91 and 1996-97. Around a million prime-age male workers are successfully reallocated each period, and 91 percent of them shift to nonexporters while 9 percent transition to exporters. The share of former exporter workers who are rehired at exporters (23 percent in both periods) exceeds the share of former nonexporter workers with a reallocation to exporters (7 percent in both periods). The small magnitude of transitions to exporters, with less than one in ten displaced workers moving to an exporter overall, and the rehiring bias at exporters towards former exporter workers suggest that labor shifts from nonexporters to exporters are not a major channel of worker reallocation.[20]

Labor market performance and economic outcomes. To assess the importance of labor reallocation for GDP, we measure the share of failed realloca-

[20]Reallocations within firms are minor (Muendler 2007): at the annual horizon, around two percent of prime-age male workers are reassigned to new jobs within their employing plant between 1990 and 1998, and less than one percent of the prime-age male workers are transferred between plants within their employing firm.

Table 6: Annual Transitions Across Firms

To:	Transitions 1990-91			Transitions 1996-97		
	Nonexp.	Exp.	*Total*	Nonexp.	Exp.	*Total*
From: (in millions)	(1)	(2)	(3)	(4)	(5)	(6)
Nonexporter	.816	.058	.874	.795	.060	.855
Exporter	.099	.030	.129	.106	.031	.137
Total	.915	.087	1.003	.901	.091	.992

Source: *RAIS* 1990-91 and 1996-97 (1% random sample), male workers nationwide, 25 to 64 years old; *SECEX* 1990-91 and 1996-97. Job accessions in Brazil within one year after separation. Employments are last employments of year (highest paying job if many), scaled (by 100) to population equivalents.

tions and the durations of successful reallocations over twelve months following displacement. The share of displaced workers with no reallocation for a year almost doubles from 25 to 46 percent between 1986 and 1998 (with minor variation across skill groups). We calculate the foregone share of GDP as the unrealized wage bill that the additional failures after 1990 imply, given a displaced worker's last wage. We only consider those displaced formal-sector workers as idle who typically become unemployed or move out of the labor force—a 36 percent share on average. So, we assume that the remaining 64 percent of displaced workers immediately take up an informal job or self employment and fully retain their pre-displacement earnings. This makes our estimates of foregone GDP conservative. The magnitudes are nevertheless striking. The unrealized wages implied by additional reallocation failures after 1990 amount to 1.4 percent of foregone GDP in 1992 and 2.4 percent in 1998. The increased duration of successful reallocations implies another half percent of foregone GDP in 1992 and .6 percent in 1998, measured similarly conservatively. This brings the total foregone wage bill to almost 2 percent of GDP in 1992, to 1 percent in 1994 (a year with strong GDP growth), and to 3 percent in 1998. We conclude that resource reallocation in the labor market is a foremost component of aggregate performance.

The descriptive evidence so far is based on unconditional means. The remainder of the paper subjects reallocation statistics to multivariate controls. Labor movements to the nontraded-output sector, for instance, can be partly due to an overvalued real exchange rate or a trend in services expansion. Concomitant reforms and firm and worker heterogeneity require attention. The following two sections analyze two main aspects of the observed employment shifts: the next section investigates reallocations to a work status other than formal employment, and the section after the next analyzes predictors of increased separations and reduced accessions. Increased separations fill, and reduced accessions fail to empty, the pool of labor to be reallocated and thus delay the average displaced worker awaiting formal-sector reallocation.

4 Work Status Transitions

To investigate how Brazil's trade exposure predicts transitions between work status, we estimate a multinomial logit model using *PME* household survey data for individual workers.[21] *PME* reports a single work status transition for every identified household member at the annual horizon.

Denote the set of work status types with $\mathbb{S}$. An individual household member's probability to move to work status $\sigma_{i,t+1}$, conditional on present work status $\sigma_{i,t} = \sigma$, is specified as

$$Pr(\sigma_{i,t+1} | \, \sigma_{i,t} = \sigma; \mathbf{x}, \mathbf{z}) = \frac{\exp\{\mathbf{z}_{S(i),t}\beta_z^\sigma + \mathbf{x}_{i,t}\beta_x^\sigma + \alpha_t^\sigma + \alpha_{c(i),t}^\sigma\}}{\sum_{\varsigma \in \mathbb{S}} \exp\{\mathbf{z}_{S(i),t}\beta_z^\varsigma + \mathbf{x}_{i,t}\beta_x^\varsigma + \alpha_t^\varsigma + \alpha_{c(i),t}^\varsigma\}}, \tag{1}$$

where $\mathbf{z}_{S(i),t}$ is a vector of sector-level covariates of the household member's initial sector $S(i)$; $\mathbf{x}_{it}$ is a vector of covariates that are job and worker specific; β_x^ς and β_z^ς are coefficient vectors for the future work status $\varsigma \in \mathbb{S}$; and α_t^ς and $\alpha_{c(i),t}^\varsigma$ are year and city effects. Coefficients are identified relative to a baseline work status at t+1. We use as the baseline work status a household member's continuation in the present work status, $\sigma_{i,t+1} = \sigma_{i,t} = \sigma$. The employer-employee specific errors of work status outcomes are assumed to be doubly exponentially distributed for the multinomial logit model and independent across employer-employee matches. For independence of the work status error to be plausible, it is important to condition on turnover characteristics of the household member. We therefore include in the vector of job-worker covariates $\mathbf{x}_{it}$ an indicator whether the household member had the same work status during the preceding four months. The employer is not identified in household data. We fit model (1) with maximum likelihood and restrict the estimation sample to manufacturing jobs at t, for which trade-related covariates $\mathbf{z}_{S(i),t}$ are well defined, but do not impose a sector restriction on job observations at t+1.

[21] We choose a multinomial over an ordered logit model because, conditional on a set of individual job and worker characteristics, work status types such as informal or self employment have no obvious intrinsic ordering.

Table 7: Work Status Transitions from Formal Employment

	From formal manufacturing employment in t to:			
(in t+1)	Informal	Self employed	Unemployed	Withdrawn
Covariate (in t)	(1)	(2)	(3)	(4)
Product Market Tariff	-1.431	-.828	.223	.490
	(.156)***	(.169)***	(.192)	(.189)***
Intm. Input Tariff	.298	.913	-1.130	-.045
	(.398)	(.436)**	(.489)**	(.495)
Formal empl. for four months	-1.767	-1.428	-.597	-1.097
	(.030)***	(.036)***	(.055)***	(.045)***
Pot. labor force experience	.005	.029	-.021	.006
	(.006)	(.008)***	(.010)**	(.008)
Some High School	.039	-.447	-.270	.295
	(.036)	(.041)***	(.051)***	(.051)***
Some College	-.038	-.827	-.734	.404
	(.086)	(.121)***	(.140)***	(.130)***
College Degree	.258	-.686	-1.151	.178
	(.050)***	(.070)***	(.107)***	(.085)**
Obs.		75,377		
Pseudo R^2		.06		

Source: *PME* 1986-99, male household members in metropolitan area, 25 years or older, with initial formal manufacturing employment (annual transitions between 4th and 8th interview). Reference category: continuation in formal work status. Controlling for year and city effects. Robust standard errors in parentheses: * significance at ten, ** five, *** one percent.

Transitions out of formality. The set of work status outcomes for a worker with a formal manufacturing job contains five alternatives: (1) the worker retains the formal manufacturing job or switches to a new formal job (not necessarily in manufacturing); (2) the worker moves to an informal job (not necessarily in manufacturing); (3) the worker moves to self-employment (not necessarily in manufacturing); (4) the worker moves to unemployment; and (5) the worker withdraws from the labor force. To capture the effect of changing tariff dispersions for a sector's effective rate of protection, we include product tariffs and intermediate-input tariffs in the multinomial logit regressions. Table 7 presents the predictions.

Reduced product-market tariffs are associated with significantly higher odds of transitions out of a formal manufacturing job and into informality or self employment. The product-market tariff is not a statistically significant predictor of unemployment but is associated with significantly higher odds of labor-force withdrawals. Intermediate-input tariff coefficients show converse signs, and predict significantly more transitions out of formality and into informality. The sign reversals are consistent with the notion of effective protection by which elevated intermediate input tariffs aggravate competitive pressure, whereas high product-tariff barriers reduce competitive pressure. As described above, Brazil's manufacturing industries face a drop in the effective rate of protection, that is a faster decline in product tariffs than in input tariffs. Together, coefficient magnitudes and the relatively faster drop of product-market tariffs imply that Brazil's trade liberalization predicts more transitions out of formal manufacturing employment and into informality or self-employment. This worker-level evidence on transitions into informality challenges findings for Brazil in Goldberg and Pavcnik (2003), who do not detect a significant effect of trade liberalization on the incidence of informality in sector data for Brazil.

Workers with continuously reported formal-sector employment during the first four months of observation are significantly less likely to lose formality status over the following year. Longer labor force experience predicts more transitions into self-employment and fewer into unemployment. Higher educational attainment predicts significantly less transitions into self employment and unemployment. But for college-educated workers the odds of a transition from formality to informality are relatively higher than for other education groups, all else equal.

Transitions out of informality. For a worker with an informal manufacturing job in the base year, alternative (1) becomes that the worker retains the job or moves to an informal job (not necessarily in manufacturing), and (2) that the worker transitions to a formal job (not necessarily in manufacturing). The remaining three work status types are as before.

Elevated product-market tariffs are associated with significantly higher odds of transitions out of informality in manufacturing and into formality or self employment. Lower product-market tariffs are also associated with significantly lower odds of a transition from informality into unemployment or withdrawals from the labor force. Intermediate-input tariff coefficients, however, are not different from zero at common significance levels. Workers with continuously reported informal-sector employment during the first four months of observation are significantly less likely to leave informality status over the following year. As for household members with initial formal work, longer labor force experience predicts more transitions into self-employment and fewer into unemployment. Higher educational attainment predicts significantly less transitions out of informality and into formal work, self employment or unemployment but has no significant effect on

Table 8: WORK STATUS TRANSITIONS FROM INFORMAL EMPLOYMENT

	From informal manufacturing employment in t to:			
(in t+1)	Formal	Self employed	Unemployed	Withdrawn
Covariate (in t)	(1)	(2)	(3)	(4)
Product Market Tariff	1.437	.735	2.141	.948
	(.255)***	(.319)**	(.614)***	(.429)**
Intm. Input Tariff	-.699	1.259	-.385	.120
	(.680)	(.816)	(1.606)	(1.124)
Informal empl. for four months	-1.323	-1.591	-1.457	-1.112
	(.037)***	(.048)***	(.106)***	(.063)***
Pot. labor force experience	-.005	.020	-.069	.006
	(.007)	(.009)**	(.019)***	(.011)
Some High School	-.377	-.667	-.725	.029
	(.042)***	(.049)***	(.106)***	(.073)
Some College	-.463	-1.131	-1.063	.028
	(.092)***	(.136)***	(.257)***	(.179)
College Degree	-.475	-1.248	-1.704	-.019
	(.058)***	(.083)***	(.211)***	(.104)
Obs.	22,246			
Pseudo R^2	.08			

Source: *PME* 1986-99, male household members in metropolitan area, 25 years or older, with initial informal manufacturing employment (annual transitions between 4th and 8th interview). Reference category: continuation in informal status. Controlling for year and city effects. Robust standard errors in parentheses: * significance at ten, ** five, *** one percent.

labor force withdrawals.

Overall, the evidence on work status transitions predicts that Brazil's tariff reductions in the 1990s are associated with significantly more moves from formal manufacturing work into informality and, at the reverse margin, with significantly lower odds that workers move from informality into formal employment.

5 Formal-Sector Separations and Accessions

To understand labor-market adjustment in the formal sector in more detail, we turn to separations and accessions and how they relate to industry, plant, job and worker characteristics. Wage-taking employers adjust their workforces through worker separations and accessions. Separations in turn burden, and accessions unburden, the pool of workers to be reallocated. So, the chance of a displaced worker to be successfully reallocated changes as economic conditions alter separation and accession rates.

Consider the probability that an employer-employee match is terminated (a separation) or is formed (an accession), conditional on a worker-fixed component α_i that is observable to the employer and the worker:

$$Pr\left(\sigma_{i,t} \middle| \mathbf{x}_{i,t}, \mathbf{y}_{J(i),t}, \mathbf{z}_{S(J(i)),t}\right) = \frac{\exp\{\mathbf{z}_{S(J(i)),t}\beta_z + \mathbf{y}_{J(i),t}\beta_y + \mathbf{x}_{i,t}\beta_x + \alpha_i + \alpha_t\}}{1 + \exp\{\mathbf{z}_{S(J(i)),t}\beta_z + \mathbf{y}_{J(i),t}\beta_y + \mathbf{x}_{i,t}\beta_x + \alpha_i + \alpha_t\}}, \tag{2}$$

where $\sigma_{i,t}$ denotes the binary outcome (accession or not, separation or not) for worker i at time t. $\mathbf{z}_{S(J(i)),t}$ is a vector of sector-level covariates of the worker's displacing or hiring sector $S(J(i))$; $\mathbf{y}_{J(i),t}$ is a vector of plant-level covariates of worker i's displacing or hiring plant $J(i)$; $\mathbf{x}_{it}$ is a vector of covariates that are worker, job or match specific; β_z, β_y, β_x are coefficient vectors; α_i is the worker-fixed effect and α_t a year effect. There is an unobserved error to terminations and formations of employer-employee matches. The error is assumed to be logistic and independent across employer-employee matches conditional on the observed covariates and the worker and year effect. We fit this conditional logit model (2) using conditional maximum likelihood estimation (the full maximum likelihood estimator is inconsistent). Identification of worker-fixed effects requires restriction of the sample to workers who experience at least one separation or accession. Coefficients on worker and job covariates are identified from time variation within and across employers. Educational attainment changes little among prime-age males, however. We consequently drop education categories from the worker characteristics vector but keep educational workforce composition shares among the plant-level regressors.

Table 9 presents conditional logit estimates of separations from formal manufacturing jobs, controlling for worker-fixed separation effects. Separations are significantly more frequent in sectors with a stronger comparative advantage and at exporters—contrary to predictions of standard trade theory. Elevated product tariffs predict lower separation rates from formal jobs (though only significant at the ten-percent level), but high input tariff barriers are associated with significantly higher separation rates. Note that high input tariffs reduce a plant's effective protection from foreign competition. Similarly, additional import penetration predicts significantly higher displacement odds. We include observed market penetration with imports to proxy for changing non-tariff barriers. Point estimates and statistical significance of these coefficients are hardly affected as the specification is gradually enriched (moving from column 1 to column 5). FDI inflows into the sector predict a statistically significant reduction in displacement

Table 9: Conditional Logit Estimation of Separations

	(1)	(2)	(3)	(4)	(5)	(6)
Balassa Comp. Adv.	.080				.169	.204
	(.021)***				(.024)***	(.023)***
Exporter Status		.289			.283	.301
		(.028)***			(.028)***	(.028)***
Product Market Tariff			-.104		-.705	-1.383
			(.416)		(.426)*	(.410)***
Intm. Input Tariff			1.601		2.880	-1.420
			(.633)**		(.678)***	(.553)**
Import Penetration				.774	1.257	6.035
				(.353)**	(.388)***	(.349)***
Sector-level covariates						
Sector real exch. rate	.733	.843	.353	.701	-.398	.213
	(.624)	(.626)	(.640)	(.631)	(.645)	(.069)***
FDI Flow (USD billion)	-.025	-.012	-.018	-.014	-.048	.047
	(.020)	(.020)	(.020)	(.020)	(.020)**	(.019)**
Herfindahl Index (sales)	-.371	-.517	-.399	-.656	-.354	.929
	(.317)	(.316)	(.329)	(.325)**	(.343)	(.320)***
Plant-level covariates						
Log Employment	-.343	-.370	-.341	-.339	-.377	-.410
	(.011)***	(.011)***	(.011)***	(.011)***	(.011)***	(.011)***
Share: Middle School or less	-.750	-.658	-.719	-.717	-.663	-.793
	(.131)***	(.131)***	(.131)***	(.131)***	(.132)***	(.129)***
Share: Some High School	-.444	-.392	-.440	-.443	-.393	-.214
	(.148)***	(.148)***	(.147)***	(.147)***	(.148)***	(.145)
Share: White-collar occ.	.721	.700	.739	.738	.691	.552
	(.075)***	(.074)***	(.074)***	(.074)***	(.075)***	(.073)***
Worker-level covariates						
Tenure at plant (in years)	1.367	1.350	1.362	1.363	1.351	1.390
	(.036)***	(.036)***	(.036)***	(.036)***	(.036)***	(.037)***
Pot. labor force experience	.006	.006	.006	.006	.006	.031
	(.002)**	(.002)**	(.002)**	(.002)**	(.002)**	(.002)***
Unskilled Wh. Collar Occ.	-.256	-.251	-.259	-.255	-.262	-.199
	(.067)***	(.067)***	(.067)***	(.067)***	(.067)***	(.065)***
Year effects	yes	yes	yes	yes	yes	no
Obs.	145,408	145,408	145,408	145,408	145,408	145,408
Pseudo R^2	.148	.149	.148	.148	.150	.137

Source: *RAIS* 1990-98 (1% random sample), male workers nationwide, 25 to 64 years old, with manufacturing job. Separations exclude transfers, deaths, and retirements. Reference observations are employments with no reported separation in a given year. Sector information at subsector IBGE level. Professional or managerial occupations and skilled blue collar occupations (not reported) not statistically significant at five-percent level. Robust standard errors in parentheses: * significance at ten, ** five, *** one percent.

rates. The sectoral real exchange and the Herfindahl concentration index have no significant predictive power after conditioning on year effects.

When we exclude year indicators from the regression (column 6), comparative advantage and exporting status become even stronger predictors of displacements. Tariffs and import penetration coefficients now also reflect the effect of reducing trade barriers over time and unambiguously predict that reduced barriers both at the input and the output margin, and the arrival of additional imports, are associated with more worker separations. Using further controls—such as the inflation rate in addition to sectoral price levels behind the real exchange rate, FDI stocks in addition to FDI flows, and controls for privatization and outsourcing—beyond the large set of sector- and firm-level variables that already control for time-varying changes to the competitive environment does not change coefficients in important ways.

Before discussing plant and worker-level variables, we turn to the opposite margin: Table 10 presents conditional logit estimates of accessions into formal manufacturing jobs, controlling for worker-fixed accession effects. Mirroring the signs from separation regressions, accession rates are lower in sectors with stronger comparative advantage, when we condition on other trade-related variables (column 5). The coefficient is not statistically significant at conventional levels in this regression (but will become statistically significant when controlling for higher-order interactions between trade variables in Table 14). Exporters exhibit significantly lower accession rates, mirroring their higher separation rates. Elevated product tariffs predict significantly more accessions, mirroring the sign from separation regression, whereas higher intermediate-input tariffs predict significantly fewer accessions, also mirroring the sign from separation regression. Import penetration has no statistically significant effect, and neither does the real exchange rate. FDI inflows are associated with significantly more accessions and more concentrated manufacturing industries exhibit fewer accessions.

When we do not condition on year effects (column 6), comparative advantage and exporting status become even stronger predictors of reduced accessions. Tariffs and import penetration coefficients now also reflect the effect of reducing trade barriers over time. Reduced input tariffs, which tend to make competition less fierce, predict more accessions. Reduced output tariffs and the arrival of additional imports, which tend to make competition more fierce, are associated with fewer accessions.

Larger manufacturing plants exhibit less turnover: they displace significantly fewer (Table 9) and they hire significantly fewer workers (Table 10). Plants with less educated workforces and more blue-collar jobs separate from workers signifi-

Table 10: Conditional Logit Estimation of Accessions

	(1)	(2)	(3)	(4)	(5)	(6)
Balassa Comp. Adv.	.041				-.016	-.114
	(.017)**				(.020)	(.019)***
Exporter Status		-.449			-.439	-.429
		(.027)***			(.027)***	(.026)***
Product Market Tariff			1.306		1.246	2.474
			(.379)***		(.393)***	(.379)***
Intm. Input Tariff			-3.258		-3.073	-3.846
			(.540)***		(.598)***	(.514)***
Import Penetration				-.522	.198	-3.919
				(.320)	(.355)	(.307)***
Sector-level covariates						
Sector real exch. rate	-1.264	-.955	-.953	-.986	-.810	.038
	(.605)**	(.606)	(.626)	(.611)	(.639)	(.076)
FDI Flow (USD billion)	.039	.047	.056	.047	.058	.031
	(.022)*	(.021)**	(.021)***	(.021)**	(.022)***	(.021)
Herfindahl Index (sales)	-.348	-.344	-.795	-.275	-.788	-2.335
	(.268)	(.268)	(.282)***	(.277)	(.297)***	(.277)***
Plant-level covariates						
Log Employment	-.190	-.140	-.189	-.189	-.141	-.112
	(.008)***	(.009)***	(.008)***	(.008)***	(.009)***	(.008)***
Share: Middle School or less	.947	.857	.940	.948	.850	.828
	(.107)***	(.105)***	(.107)***	(.107)***	(.105)***	(.104)***
Share: Some High School	.740	.667	.739	.740	.668	.468
	(.124)***	(.122)***	(.124)***	(.124)***	(.122)***	(.120)***
Share: White-collar occ.	-.675	-.614	-.679	-.671	-.621	-.534
	(.067)***	(.067)***	(.067)***	(.067)***	(.067)***	(.064)***
Worker-level covariates						
Prof. or Manag'l. Occ.	-.801	-.807	-.801	-.800	-.807	-.827
	(.068)***	(.068)***	(.068)***	(.068)***	(.068)***	(.066)***
Tech'l. or Superv. Occ.	-.603	-.610	-.597	-.603	-.604	-.623
	(.064)***	(.064)***	(.064)***	(.064)***	(.064)***	(.062)***
Unskilled Wh. Collar Occ.	-.490	-.497	-.488	-.489	-.495	-.519
	(.061)***	(.062)***	(.062)***	(.061)***	(.062)***	(.060)***
Skilled Bl. Collar Occ.	-.417	-.413	-.413	-.417	-.410	-.443
	(.032)***	(.032)***	(.032)***	(.032)***	(.032)***	(.031)***
Year effects	yes	yes	yes	yes	yes	no
Obs.	112,974	112,974	112,974	112,974	112,974	112,974
Pseudo R^2	.036	.040	.037	.036	.041	.026

Source: *RAIS* 1990-98 (1% random sample), male workers nationwide, 25 to 64 years old, with manufacturing job. Accessions exclude transfers. Reference observations are employments with no reported accession. Sector information at subsector IBGE level. Robust standard errors in parentheses: * significance at ten, ** five, *** one percent.

cantly less frequently and hire significantly more frequently. Workers with a longer tenure at the plant and longer labor-market experience suffer significantly more frequent separations at the separation margin. Workers in occupations of intermediate skill intensity experience significantly fewer separations, and workers are significantly less likely to be hired into high-skill intensive manufacturing occupations (with a monotonic drop in accession odds as an occupation's skill intensity increases). Year effects are significant at the one-percent level and show both a strictly monotonic increase in manufacturing separations and a strictly monotonic drop in manufacturing accessions (see Table 16 in the Appendix).

Worker heterogeneity is an important predictive component of separations and accessions. A comparison between conditional and unconditional logit estimation (not reported here) shows that regressions are highly sensitive to the omission of worker-fixed effects. The relevance of conditional worker effects is consistent with the hypothesis that the termination and formation of employer-employee matches is not random, even after controlling for a comprehensive set of observable worker and employer characteristics.

To gain a sense of how important trade is for labor market outcomes in Brazil's manufacturing sector, we use changes in import penetration rates and tariffs since 1990 to predict changes in separation and accession rates, based on conditional logit estimates from Tables 9 and 10 (columns 6). Import penetration more than doubles between 1990 and 1998, while product-market and input tariffs drop by more than half. Additional import penetration in 1998, beyond its 1990 level, predicts a 6-percent increase in the separation rate by 1998. Tariff reductions below 1990 levels predict an 8-percent increase in the separation rate. These are salient magnitudes compared to the observed separation rate fluctuation over the 1990s (see Figure 2 above). The manufacturing sector employs roughly one in four male workers in the national labor force so that these predictions also matter for the aggregate. At the accession margin, additional import penetration predicts a 4-percent reduction in hiring rates, whereas the ambiguous effect of product and input tariffs on accession rates partly counteracts the prediction. Overall, the magnitudes suggest that trade is a potentially important source of national labor-market performance. However, the predictions do not reflect additional firm-level responses to trade, including productivity change and export-status transitions, or implied labor shifts across industries. A comprehensive evaluation arguably calls for a structural trade model. It is thus tantamount to assess whether known trade models adequately capture the direction of labor flows across employers and industries.

Table 11: LINEAR AND INSTRUMENTAL-VARIABLE ESTIMATION

	Separations			Accessions		
		OLS-FE			OLS-FE	
	Cdl. logit		IV	Cdl. logit		IV
	(1)	(2)	(3)	(4)	(5)	(6)
Balassa Comp. Adv.	.169	.017	.023	-.016	.002	-.002
	(.024)***	(.002)***	(.003)***	(.020)	(.002)	(.003)
Exporter Status	.283	.038	.516	-.439	-.049	-.500
	(.028)***	(.003)***	(.096)***	(.027)***	(.003)***	(.091)***
Product Market Tariff	-.705	-.100	-.032	1.246	.124	.113
	(.426)*	(.035)***	(.081)	(.393)***	(.032)***	(.073)
Intm. Input Tariff	2.880	.343	.161	-3.073	-.309	-.227
	(.678)***	(.054)***	(.141)	(.598)***	(.049)***	(.132)*
Import Penetration	1.257	.052	.004	.198	.088	.265
	(.388)***	(.034)	(.077)	(.355)	(.031)***	(.071)***
Obs.	145,408	293,353	293,353	112,974	293,124	293,124

Source: *RAIS* 1990-98 (1% random sample), male workers nationwide, 25 to 64 years old, with manufacturing job. Separations and accessions exclude transfers, deaths, and retirements. Reference observations are employments with no reported separation or accession in a given year. Sector information at subsector IBGE level. Estimates in column 1 and 4 repeat column 6 in Tables 9 and 10. Further regressors (not reported): Year indicators, sector, plant and worker covariates. Robust standard errors in parentheses: * significance at ten, ** five, *** one percent.

6 Concomitant Economic Changes and Reforms

Brazil's trade liberalization predicts changes in worker separation and accession rates across sectors. But neither comparative-advantage sectors nor exporters exhibit the expected labor absorption; they separate from their workers significantly more frequently than other sectors and firms. Exporters also hire significantly less frequently.[22] We address potential empirical concerns for these predictions of worker flows. We consider the potential simultaneity of trade policies and exporting status, the relevance of Brazil's concomitant reforms and the role of firm-level labor productivity.

[22]Direct estimation of reallocation durations (using rehiring hazards for prime-age male workers after separation from a formal-sector manufacturing job) corroborate the evidence on separation and accession rates. We report the duration estimates in our working paper (Menezes Filho and Muendler 2007).

Trade exposure and exporting status. Despite the apparently exogenous nature of trade reform for individual employers—the enactment by decree on president Collor's first day in office surprises politicians and businesses alike—, the reduction in tariff dispersion gives rise to a simultaneity concern. By design, initially highly protected sectors face the largest product tariff declines. Similarly, market penetration with foreign inputs possibly responds to Brazilian labor-market conditions. We therefore predict tariffs and market penetration rates at the sector level with instrumental variables. At the firm-level, employers decide exporting status and labor turnover simultaneously. We therefore also predict export status with instrumental variables.

To construct instruments for export demand, we consider seven broad destination regions of Brazil's exports and calculate the destination-region imports from other source-countries than Brazil. These foreign demand proxies vary by sector and year. In addition, we employ the nominal U.S. dollar exchange rate and sector-level producer-price indices in the United States and the EU as instruments. Nominal exchange rate movements are largely unpredictable, and foreign producer prices in industrialized economies are arguably exogenous to Brazil. To check for potential sign reversals and assess the magnitude of possible simultaneity bias, we resort to linear fixed-effects regressions of separation and accession indicators on the same predictors as in the preceding section:[23]

$$\sigma_{i,t} = \hat{\mathbf{z}}_{S(J(i)),t}\beta_z + \hat{\mathbf{y}}_{J(i),t}\beta_y + \mathbf{x}_{i,t}\beta_x + \alpha_i + \alpha_t + \epsilon_{i,t}, \tag{3}$$

where $\sigma_{i,t} \in \{0, 1\}$ denotes the binary outcome (accession or not, separation or not) for worker i at time t, and regressor and coefficient vectors are as in (2). There is an unobserved error $\epsilon_{i,t}$ to the termination and formation of employer-employee matches. It is assumed to be normally distributed and independent across employer-employee matches. We first predict the subset of potentially simultaneity-afflicted regressors in $\mathbf{z}_{S(J(i)),t}$ and $\mathbf{y}_{J(i),t}$ with instrumental variables, and then include their predictions $\hat{\mathbf{z}}_{S(J(i)),t}$ and $\hat{\mathbf{y}}_{J(i),t}$ in (3). Turning to linear regression has the additional benefit that the estimation sample includes workers with no change in employment; their worker-fixed effect is separately identified through time variation of other predictors at the same employer. The change in estimation sample affords an additional robustness check.

[23]Linearly predicting export status, product and input tariffs, and import penetration with the instruments, and including both predicted values and residuals in conditional logit estimation, shows coefficients on the residuals to be statistically significant and renders simultaneity a potential empirical issue (Rivers and Vuong 1988).

On the first stage, we regress export status, product and input tariffs, and import penetration on the instrumental variables, weighting the regression by employment observations in the separation and accession samples. Table 17 in the Appendix shows results by sample (except for the input-tariff estimates which are similar to product-tariff estimates). There is no evidence of weak instruments: F statistics from joint significance tests on the instruments vary between 13 and 14,000. Almost invariably, the instruments are statistically significant predictors at the one-percent level.[24] We highlight a few coefficient estimates. Expectedly, higher producer prices in the United States and Europe, as well as a weaker Brazilian currency, predict significantly more frequent exporting status. Employment-weighted exporting status is more frequent in sectors with weaker comparative advantage, as documented in the data Section 2 before, because there is a larger number of small-volume exporters in the low-advantage sectors.

Table 11 redisplays conditional logit estimates for separations and accessions in columns 1 and 4 (from Tables 9 and 10, column 6). We compare those estimates to linear worker-fixed effects regressions without (columns 2 and 5) and with instrumental-variable predictions (columns 3 and 6). The estimation samples for the linear worker-fixed effects models are substantially larger because workers with no transition remain in the sample. When instrumenting, there is not a single sign reversal in the potentially simultaneity-afflicted coefficients—export status, tariffs, and import penetration (comparing columns 2 and 3, and columns 5 and 6). Instrumentation overwhelmingly reinforces at the one-percent significance level that comparative-advantage sectors and exporters exhibit more separations, and exporters exhibit fewer accessions. Several coefficients on tariffs and import penetration lose significance at common levels under instrumental-variable fixed-effects regressions (columns 3 and 6) but never exhibit a sign reversal. So, instrumentation in a linear probability model corroborates our main explanation for lacking labor reallocation: firms in comparative-advantage sectors and exporters separate from their workers significantly more frequently than the average employer, and exporters hire significantly less frequently.

Economic change and reforms. The Brazilian economy undergoes a series of concomitant economic transformations during the sample period, including technological changes, the intensified outsourcing of service jobs, surging foreign direct

[24] We also experiment with labor productivity in the initial year 1990 as a candidate firm-level instrument in the subsample of *PIA* firms but over-identification tests reject its validity when added.

Table 12: Alternative Specifications

	Cdl. logit	Young worker	Primary school	College educ.	Sector FE	Privat. control	Outsrc. job ind.
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
				Separations			
Balassa Cmp. Adv.	.169	.498	.145	.216	-.094	.170	.169
	(.024)***	(.267)*	(.028)***	(.150)	(.049)*	(.026)***	(.024)***
Exporter Status	.283	.379	.296	.297	.284	.283	.283
	(.028)***	(.243)	(.033)***	(.143)**	(.028)***	(.028)***	(.029)***
Product Mkt. Trff.	-.705	-3.960	-.500	-1.771	-2.361	-.694	-.751
	(.426)*	(4.290)	(.499)	(2.281)	(.476)***	(.427)	(.430)*
Intm. Input Trff.	2.880	10.027	2.469	7.146	5.149	2.875	3.010
	(.678)***	(7.163)	(.779)***	(4.086)*	(.748)***	(.675)***	(.686)***
Import Penetration	1.257	8.588	.678	.886	3.227	1.264	1.269
	(.388)***	(3.668)**	(.477)	(1.995)	(.638)***	(.392)***	(.391)***
addl. regressor(s)					yes	-.142	-.018
						(1.227)	(.037)
Obs.	145,408	2,897	110,831	7,498	145,408	145,408	143,536
Pseudo R^2	.150	.391	.161	.245	.151	.150	.151
				Accessions			
Balassa Cmp. Adv.	-.016	-.120	-.006	-.141	-.067	-.024	-.015
	(.020)	(.209)	(.023)	(.118)	(.048)	(.022)	(.021)
Exporter Status	-.439	-.477	-.420	-.776	-.438	-.439	-.437
	(.027)***	(.216)**	(.031)***	(.140)***	(.027)***	(.027)***	(.027)***
Product Mkt. Trff.	1.246	.099	1.333	2.033	1.822	1.118	1.185
	(.393)***	(3.290)	(.451)***	(2.092)	(.498)***	(.412)***	(.397)***
Intm. Input Trff.	-3.073	-7.113	-2.943	-5.152	-2.954	-2.987	-3.041
	(.598)***	(5.668)	(.673)***	(3.393)	(.750)***	(.603)***	(.604)***
Import Penetration	.198	-9.315	.084	-.720	1.764	.128	.181
	(.355)	(3.845)**	(.423)	(1.948)	(.665)***	(.363)	(.358)
addl. regressor(s)					yes	1.140	-.098
						(1.166)	(.033)***
Obs.	112,974	2,752	86,468	4,786	112,974	112,974	110,985
Pseudo R^2	.041	.223	.043	.088	.042	.041	.040

Source: *RAIS* 1990-98 (1% random sample), male workers nationwide, 25 to 64 years old, with manufacturing job. Estimates in column 1 repeat column 6 in Tables 9 and 10. Further regressors (not reported): Year indicators, sector, plant and worker covariates. Robust standard errors in parentheses: * significance at ten, ** five, *** one percent.

investment inflows and policy shifts such as macroeconomic stabilization, capital-account liberalization, and privatization. Labor-market institutions were altered preceding trade reform. The accession and separation regressions so far control for sector and year covariates including sectoral real exchange rates, Herfindahl sales concentration indices, foreign direct investment inflows, and year fixed effects. We turn to economic changes and policies that perhaps affect estimates at the level of the plant, job, worker or employer-employee match in specification (2).

If skill-biased technological change systematically interacts with the effect of trade reform on labor turnover, trade reform expectedly covaries with labor turnover differently for workers with different skills. We run specification (2) separately for young workers with less than ten years of potential labor-market experience, and for workers with primary schooling and some college education. Table 12 redisplays in column 1 the conditional logit estimates for separations and accessions on the full sample. Estimates for the skill subsamples follow in columns 2 through 4. Coefficient estimates for separations and accessions are strikingly similar across the samples. No sign changes. Import penetration predicts a stronger effect on young workers' separations and accessions and, surprisingly, implies that trade integration predicts more frequent separations and significantly less frequent accessions for young workers. This also suggests that, if anything, our restriction of the regression sample to prime-age workers biases trade effects against us. Magnitudes of the tariff and import-penetration coefficients significantly increase for more educated workers, but are statistically indistinguishable for comparative advantage and export status. Statistical significance is lost in some cases in the smaller college-educated worker subsample. There is, to our surprise, no strong evidence that skill-biased labor-demand changes systematically interact with the effect of trade reform on separations and accessions.

Though the constitutional labor-market reforms in 1988 precede trade reform in 1990, they might affect sectors with unobserved differences in workforce composition to varying degrees and interact with trade reform in a way that erroneously attributes labor turnover to the trade regime. We use sector-fixed effects at the subsector IBGE level to capture unobserved sectoral differences in the effect of labor institutions on unobserved separation and accession determinants. Table 12 reports estimates in column 5. Expectedly, inclusion of sector indicators turns the coefficient on comparative advantage, which is highly sector specific and largely time invariant in our data, insignificant. For the other trade regressors, however, coefficient estimates increase in absolute value (compared to column 1) and remain highly significant. These findings, and the evidence from instrumental-variable regressions, render erroneous attribution of labor-market regulation effects to trade reform little plausible.

The privatization of state-owned businesses and the progressing outsourcing of service jobs to specialized suppliers can affect separations and accessions. If privatization and outsourcing covary with the trade regime and labor turnover in systematic ways, they potentially lead to erroneous attributions. The ownership status of a plant is observable in *RAIS* since 1995, when the federal government started to pursue privatization on a larger scale. We impute a plant's ownership status in 1990-94 as the ownership status in 1995 and include the private-ownership indicator at the plant-level in regression (2). As column 6 in Table 12 shows, coefficient estimates on the trade-related variables exhibit no statistically significant change, and the ownership-status itself is not a statistically significant predictor. We infer the susceptibility of a job to outsourcing (*tercerização*) if it is a service occupation at the CBO three-digit level that can be performed in-house or be provided by a specialized subcontractor. Including the job-level indicator in regression (2) results in no statistically significant coefficient change (column 7). Jobs susceptible to outsourcing exhibit a statistically significant reduction in accession odds. There is, in summary, no evidence that simultaneous economic changes and concomitant reforms systematically alter the effect of trade reform on separations and accessions.

Firm-level labor productivity. Exporters are more productive than nonexporters (Table 3). To compare the relative importance of a firm's exporter status and labor productivity for separations and accessions, we include a measure of firm-level labor productivity in specification (2). For this purpose, we use the subsample of *RAIS* firms that are surveyed in *PIA*, for which firm-level labor productivity is inferrable. Table 11 redisplays conditional logit estimates for separations and accessions on the full sample in columns 1 and 4 (from Tables 9 and 10, column 6). The table compares those prior estimates to estimates on the combined *PIA*-*RAIS* subsample without (columns 2 and 5) and with log labor productivity as a regressor (columns 3 and 6). The export-status coefficient loses statistical significance in the reduced separation subsample of *PIA* manufacturers but does not change sign. Exporters exhibit significantly fewer accessions at the one-percent level; this reinforces our prior finding. Trade-variables, including Balassa comparative advantage, are not significant predictors of separations and accessions in the reduced subsample. Higher labor productivity itself predicts significantly fewer accessions. This is consistent with the descriptive evidence (Table 2) that faster labor productivity growth at manufacturing firms correlates

Table 13: CONDITIONAL LOGIT ESTIMATION WITH LABOR PRODUCTIVITY

	Separations			Accessions		
	Full smpl.	*PIA* smpl.		Full smpl.	*PIA* smpl.	
	(1)	(2)	(3)	(4)	(5)	(6)
Balassa Comp. Adv.	.169	-.006	-.006	-.016	-.017	-.012
	(.024)***	(.074)	(.074)	(.020)	(.060)	(.060)
Exporter Status	.283	.030	.030	-.439	-.291	-.286
	(.028)***	(.076)	(.076)	(.027)***	(.075)***	(.075)***
Product Market Tariff	-.705	1.245	1.244	1.246	-.337	-.262
	(.426)*	(.987)	(.987)	(.393)***	(.955)	(.955)
Intm. Input Tariff	2.880	.373	.380	-3.073	-1.241	-1.531
	(.678)***	(1.608)	(1.611)	(.598)***	(1.374)	(1.378)
Import Penetration	1.257	1.332	1.327	.198	.460	.625
	(.388)***	(1.000)	(1.004)	(.355)	(1.098)	(1.099)
Log Labor Productivity			.003			-.115
			(.051)			(.053)**
Obs.	145,408	40,335	40,335	112,974	20,191	20,191
Pseudo R^2	.150	.335	.335	.041	.089	.089

Source: *RAIS* 1990-98 (1% random sample) linked to *PIA* 1990-98, male workers nationwide, 25 to 64 years old, with manufacturing job. Separations and accessions exclude transfers, deaths, and retirements. Reference observations are employments with no reported separation or accession in a given year. Sector information at subsector IBGE level. Further regressors (not reported): Year indicators, sector, plant and worker covariates. Robust standard errors in parentheses: * significance at ten, ** five, *** one percent.

with slower-than-average workforce growth. Overall, the inclusion of log labor productivity in a smaller random sample of manufacturers overturns none of our results and significantly reinforces several findings.

7 Labor Market Evidence and Trade Theory

Our finding that neither comparative-advantage sectors nor exporters absorb displaced workers after trade reform challenges classic trade theory (Ricardo, Heckscher-Ohlin-Samuelson) and recent firm-level trade models (Bernard, Eaton, Jensen and Kortum 2003, Melitz 2003). Import penetration intensifies after trade reform, and significantly more workers are displaced when employers face stronger import penetration. But employers in comparative-advantage sectors and exporters separate from workers significantly more frequently, and exporters hire significantly less frequently than the average firm.

Extensions of classic trade theory recognize the potential importance of reallocation frictions. Mussa (1978), for instance, introduces adjustment costs to factor employment into the Heckscher-Ohlin-Samuelson model and shows that the long-run equilibrium critically depends on the adjustment technology.[25] Our data allow us to discern between adjustments at the separation and accession margins and suggest that the main concern is not a lacking employment reduction; worker separations significantly increase with import penetration. It is the paucity of re-accessions after separations that characterizes the failure of formal-sector reallocation.

Aspects of Brazil's experience might be perceived as consistent with predictions of recent trade models that make factor-market institutions a source of comparative advantage and find that countries with less rigid factor markets tend to specialize in industries with high factor turnover (Saint Paul 1997, Davidson, Martin and Matusz 1999, Cunat and Melitz 2006). Brazil's comparative-advantage sectors indeed exhibit more labor turnover: significantly higher worker separation rates and, unconditionally, higher worker accession rates. The lacking net expansion of comparative advantage sectors, however, is not compatible with that explanation. Moreover, comparing World Bank indices of labor-market rigidity for Brazil to weighted averages of Brazil's trading partners shows that Brazil's labor market is considerably more rigid.[26] So, those theories would predict Brazil to specialize in industries with low labor turnover, contrary to our evidence.

Bernard, Redding and Schott (2007) embed heterogeneous firms in a classic trade model and derive predictions for labor turnover. Their setting preserves the prediction from classic trade theory that there is net job creation in comparative-advantage industries and net job destruction in disadvantage industries. In the presence of productivity dispersion across firms, however, important differences between gross and net job creation and destruction result. In disadvantage industries, where there is net job destruction, high-productivity firms expand to serve the export market and create new jobs. In comparative-advantage industries, where there is net job creation, existing jobs are destroyed at low-productivity

[25]Rigid real wages, which increase throughout the 1990s in Brazil, are another known cause for hampered reallocation in trade models (Brecher 1974).

[26]For the World Bank's four rigidity and difficulty indices (hiring difficulty, hours rigidity, firing difficulty, employment rigidity) and its firing-cost measure in weekly wage equivalents, Brazil exhibits mean values between 67 and 165, whereas the mean values for Brazil's trading partners vary between 20 and 49 for three choices of trade weighting (we consider trade volume, source-country import and destination-country export weighting using *WTF* (NBER) data for Brazil). Results are available online at *wwww.econ.ucsd.edu/muendler/research*.

Table 14: CONDITIONAL LOGIT ESTIMATION WITH INTERACTIONS

	Separations			Accessions		
	(1)	(2)	(3)	(4)	(5)	(6)
Balassa Cmp. Adv.	.169	.138	.134	-.016	-.058	-.125
	(.024)***	(.036)***	(.043)***	(.020)	(.032)*	(.038)***
Cmp. Adv. × Prd. Trff.		.202	.265		.289	.599
		(.200)	(.238)		(.162)*	(.203)***
Exporter Status	.283	.481	.478	-.439	-.359	-.564
	(.028)***	(.048)***	(.081)***	(.027)***	(.045)***	(.077)***
Exporter × Prd. Trff.		-1.071	-.950		-.428	.351
		(.213)***	(.362)***		(.195)**	(.323)
Cmp. Adv. × Exporter			.011			.156
			(.051)			(.047)***
... × Prd. Trff.			-.141			-.680
			(.291)			(.250)***
Product Market Tariff	-.705	-.424	-.499	1.246	.967	.541
	(.426)*	(.532)	(.548)	(.393)***	(.474)**	(.504)
Intm. Input Tariff	2.880	3.241	3.287	-3.073	-2.486	-2.297
	(.678)***	(.767)***	(.767)***	(.598)***	(.672)***	(.682)***
Import Penetration	1.257	1.093	1.088	.198	.035	-.0008
	(.388)***	(.393)***	(.393)***	(.355)	(.364)	(.364)
Obs.	145,408	145,408	145,408	112,974	112,974	112,974
Pseudo R^2	.150	.150	.151	.041	.041	.041

Source: *RAIS* 1990-98 (1% random sample), male workers nationwide, 25 to 64 years old, with manufacturing job. Separations and accessions exclude transfers, deaths, and retirements. Reference observations are employments with no reported separation or accession in a given year. Sector information at subsector IBGE level. Further regressors (not reported): Year indicators, sector, plant and worker covariates. Columns 1 and 4 repeat column 6 of Tables 9 and 10. Robust standard errors in parentheses: * significance at ten, ** five, *** one percent.

firms.[27]

An empirical investigation of the Bernard et al. (2007) model's labor-market predictions calls for the inclusion of higher-order interactions between trade reform, comparative advantage and exporting status. Table 14 compares our previous separation and accessions estimates in columns 1 and 4 (from Tables 9 and 10, column 6) to regressions with interaction terms in the remaining columns. There are no remarkable changes to coefficient estimates for separations. At the

[27]Formally, existing jobs are destroyed at low-productivity firms that exit. But a firm exit could also be interpreted as a plant closure within a firm or as the shutdown of a product line within a plant.

accession margin, however, three noteworthy changes emerge for the full set of interactions (column 6). First, the negative comparative advantage coefficient turns significant: employers in a comparative-advantage sector hire workers significantly less frequently. So, the classic-trade prediction that there is net job creation in comparative-advantage industries is statistically significantly refuted; comparative-advantage industries separate from significantly more workers and hire significantly fewer workers. Second, product tariff reductions depress accession rates most strongly in comparative-advantage industries, conditional on exporter presence. Third, although exporters hire significantly fewer workers in general, within comparative-advantage industries exporters hire significantly more workers than nonexporters and product-tariff cuts magnify the exporter-nonexporter difference. The latter two findings are consistent with a main firm-level prediction of the Bernard et al. (2007) model: in comparative-advantage industries, existing jobs are destroyed less frequently at exporters.

None of the aforementioned explanations allows for possibly trade-induced productivity improvements within surviving firms and the labor-market consequences. Heckscher-Ohlin-Samuelson textbook models that consider sector-wide productivity change show higher productivity to reduce sector-wide employment (unless highly elastic consumer demand raises output more than proportional compared to labor productivity, e.g. Obstfeld and Rogoff 1996, 4.3.2). Recent research provides firm-level underpinnings to such sector-wide productivity effects. Raith (2003), for instance, shows in a spatial-differentiation model with free entry and exit on a unit circle that tougher product-market competition (due to closer product substitutability) induces exits, shifts product-market shares to survivors, and provides stronger managerial incentives to raise production efficiency. In recent unpublished work, Bernard, Redding and Schott (2006) attribute within-firm productivity to the composition of a firm's product range and show that increased trade exposure can raise firm-level productivity through a specialization of the firm in high-efficiency goods. If factor productivity rises faster than output in a general-equilibrium extension to the Raith (2003) model or in the Bernard et al. (2006) framework, increased trade exposure can generate Brazil's observed productivity growth in the presence of product-market share reallocations to more productive firms and labor reallocation away from more productive firms.

8 Conclusions

This paper contrasts the common finding that output shares are reallocated to more productive firms after trade reform with direct evidence on the factor market. A comprehensive linked employer-employee data set tracks formal-sector workers across employers and industries in the aftermath of Brazil's large-scale trade reform. The paper documents that comparative-advantage industries and exporters impede, rather than foster, the formal-sector labor reallocations needed to absorb workers after trade-induced displacements. Employers in comparative-advantage industries and exporters separate from significantly more workers and hire significantly fewer workers than the average firm. As a consequence, trade opening is associated with more frequent transitions to informal work status and unemployment, longer durations of formal-job reallocations and more frequent failures of formal-job reallocations for an extended period of time.

The focus on labor reallocation is not suited for a comprehensive welfare evaluation of trade opening. Gains from trade through access to more varieties of goods at undistorted relative prices accrue, even in the absence of factor reallocation. But lacking labor-market adjustment with idle resources for extended periods of time suggests that piecemeal reform can be preferable to radical policy rupture. Brazil's evidence cautions against the hypothesis that pro-competitive reform did not go far enough for economic growth to respond. To the contrary, more frequent failures of worker reallocations in the formal sector, more frequent transitions to informal work status and unemployment, more frequent withdrawals from the labor force, and longer durations of worker reallocations after large-scale trade reform burden Brazil's economic activity and are adverse to growth. Although product-market reallocation can be rapid after trade reform, countries similar to Brazil may want to prepare for prolonged and incomplete adjustment in the labor market.

Appendix

A Linked employer-employee data

Brazilian law requires every Brazilian plant to submit detailed annual reports with individual information on its workers and employees to the ministry of labor (*Ministério de Trabalho*, MTE). The collection of the reports is called *Relação Anual de Informações Sociais*, or *RAIS*, and typically concluded at the parent firm by late February or early March for the preceding year of observation. *RAIS* primarily provides information to a federal wage supplement program (*Abono Salarial*), by which every worker with formal employment during the calendar year receives the equivalent of a monthly minimum wage. A strong incentive for compliance is that workers' benefits depend on *RAIS* so that workers follow up on their records. The payment of the worker's annual public wage supplement (*Abono Salarial*) is exclusively based on *RAIS* records. The ministry of labor estimates that currently 97 percent of all formally employed workers in Brazil are covered in *RAIS*, and that coverage exceeded 90 percent throughout the 1990s.

Observation screening. In *RAIS*, workers are identified by an individual-specific PIS (*Programa de Integração Social)* number that is similar to a social security number in the United States (but the PIS number is not used for identification purposes other than the administration of the wage supplement program *Abono Salarial*). A given plant may report the same PIS number multiple times within a single year in order to help the worker withdraw deposits from the worker's severance pay savings account (*Fundo de Garantia do Tempo de Serviço*, *FGTS*) through spurious layoffs and rehires. Bad compliance may cause certain PIS numbers to be recorded incorrectly or repeatedly. To handle these issues, we screen *RAIS* in two steps. (1) Observations with PIS numbers shorter than 11 digits are removed. These may correspond to informal (illegal) workers or measurement error from faulty bookkeeping. (2) For several separation statistics, we remove multiple jobs from the sample if a worker's duplicate jobs have identical accession and separation dates at the same plant. For a worker with such multiple employments, we only keep the observation with the highest average monthly wage level (in cases of wage level ties, we drop duplicate observations randomly).

Experience, education and occupation categories. For the years 1986-93, *RAIS* reports a worker's age in terms of eight age ranges. For consistency, we categorize the age in years into those eight age ranges also for 1994-2001. We construct a proxy for potential workforce experience from the nine education categories and the mean age within a worker's age range. For example, a typical Early Career worker (34.5 years of age) who is also a Middle School Dropout (left school at 11 years of age) is assigned 23.5 years of potential workforce experience.

The following tables present age and education classifications from *RAIS*, along with the imputed ages used in construction of the potential experience variable. We use the age range information in our version of *RAIS* to infer the "typical" age of a worker in the age range as follows:

	RAIS Age Category	Imputed Age
1.	Child (10-14)	*excluded*
2.	Youth (15-17)	*excluded*
3.	Adolescent (18-24)	*excluded*
4.	Nascent Career (25-29)	27
5.	Early Career (30-39)	34.5
6.	Peak Career (40-49)	44.5
7.	Late Career (50-64)	57
8.	Post Retirement (65-)	*excluded*

For regression analysis, our education variable regroups the nine *RAIS* education categories into four categories as follows:

	Education Level	*RAIS* Education
1.	Illiterate, or Primary or Middle School Educated	1-5
2.	Some High School or High School Graduate	6-7
3.	Some College	8
4.	College Graduate	9

Occupation indicators derive from the 3-digit CBO classification codes in our nationwide *RAIS* data base, and are reclassified to conform to the ISCO-88 categories.[28] We map ISCO-88 categories to *RAIS* occupations as follows:

[28]See the online documentation at *www.econ.ucsd.edu/muendler/brazil.*

	ISCO-88 Category	Occupation Level
1.	Legislators, senior officials, and managers	Professional & Managerial
2.	Professionals	Professional & Managerial
3.	Technicians and associate professionals	Technical & Supervisory
4.	Clerks	Other White Collar
5.	Service workers and shop and market sales workers	Other White Collar
6.	Skilled agricultural and fishery workers	Skill Intensive Blue Collar
7.	Craft and related workers	Skill Intensive Blue Collar
8.	Plant and machine operators and assemblers	Skill Intensive Blue Collar
9.	Elementary occupations	Other Blue Collar

B Manufacturing firm data

For robustness checks in Table 13, we use productivity measures from Brazil's annual manufacturing firm survey *PIA* (*Pesquisa Industrial Anual*) for 1986-98. *PIA* is a representative sample of all but the smallest manufacturing firms, collected by Brazil's statistical bureau IBGE. We first obtain log TFP measures from Olley and Pakes (1996) estimation at the *Nível 50* sector level under a Cobb-Douglas specification (Muendler 2004). We then convert log TFP to log labor productivity by adding the production-coefficient weighted effects of capital accumulation and intermediate input use. Labor productivity is denominated in BRL-deflated USD-1994 output equivalents per worker.

IBGE's publication rules allow data from *PIA* to be withdrawn in the form of tabulations with at least three firms per entry. We construct random combinations of three firms by drawing from sector-location-year cells. A cell is defined by the firm's *Nível 50* sector, headquarters location, and pattern of observation years. We assign every *PIA* firm to one and only one multi-firm combination. Per cell, one four- or five-firm combination is defined when the number of firms in the sector-location-year cell is not divisible by three. For each three-to-five-firm combination, we calculate mean log productivity but retain the firm identifiers behind the combination—permitting the linking to *RAIS*.

C Additional sector data

We use data on *ad valorem* tariffs by sector and year from Kume, Piani and Souza (2000). We combine these tariff series with economy-wide input-output matrices from IBGE to arrive at intermediate input tariff measures by sector and year. We calculate the vector of sector-level input tariff indices as $\tau^{in}_{i,t} = w'_{i,t}\tau^{out}_{i,t}$ in year t, where $w_{i,t}$ is the matrix of sector-specific shares of inputs. We combine tariff with sector-average value added from *PIA* to calculate effective rates of protection by sector and year. The vector of sector-level effective rates of protection is defined as $ERP_{i,t} \equiv (\tau^{in}_{i,t} - \bar{\alpha}_{i,t}\tau^{out}_{i,t})/(1-\bar{\alpha}_{i,t})$, where $\bar{\alpha}_{i,t}$ is the sector mean of intermediate input shares in output.

We use Ramos and Zonenschain (2000) national accounting data to calculate market penetration with foreign imports. Arguably, domestic firms find the absorption market corresponding to $A_{i,t} \equiv Y_{i,t} - (X_{i,t} - M_{i,t})$ the relevant domestic environment in which they compete. We define the effective rate of market penetration as $M_{i,t}/A_{i,t}$. Foreign direct investment (FDI) and annual GDP data are from the Brazilian central bank.

We construct sector-specific real exchange rates from the nominal exchange rate to the U.S. dollar E, Brazilian wholesale price indices P_i, and average foreign price series for groups of Brazil's main trading partners P^*_i by sector i, and define the real exchange rate as $q_i \equiv EP^*_i/P_i$ so that a high value means a depreciated real sector exchange rate. We rebase the underlying price series to a value of 1 in 1995. We use Brazil's import shares from its major 25 trading partners in 1995 as weights for P^*_i. We obtain sector-specific annual series from producer price indices for the 12 OECD countries among Brazil's main 25 trading partners (sector-specific PPI series from *SourceOECD*; U.S. PPI series from *Bureau of Labor Statistics*). We combine these sector-specific price indices with the 13 annual aggregate producer (wholesale if producer unavailable) price index series for Brazil's remaining major trading partners (from *Global Financial Data*), for whom sector-specific PPI are not available.

Table 15: Subsector IBGE and *Nível 50* Comparison

Subsector IBGE	Comp. Adv.		Quintile	
Nível 50	1990	97	90	97
2 **Manufacture of non-metallic mineral products**	**.994**	**1.047**	**3**	**3**
4 Manufacture of nonmetallic mineral products	1.122	1.242	3	3
3 **Manufacture of metallic products**	**1.696**	**1.498**	**4**	**4**
5 Manufacture of iron and steel products	2.912	2.170	4	4
6 Manufacture of nonferrous metal products	1.923	1.669	4	4
7 Manufacture of metal products n.e.c.	1.426	1.267	4	3
4 **Manufacture of machinery, equipment and instruments**	**.461**	**.575**	**1**	**1**
8 Manufacture of machinery and commercial equipment	.507	.650	1	2
5 **Manufacture of electrical and telecomm. equipment**	**.523**	**.611**	**1**	**2**
10 Manufacture of electrical equipment and components	.432	.467	1	1
11 Manufacture of electronic and communication equipment	.453	.487	1	1
6 **Manufacture of transport equipment**	**1.044**	**.967**	**4**	**3**
12 Manufacture of automobiles, trucks and buses	.746	1.020	2	3
13 Manufacture of vehicle parts and transportation eqpmt.	.802	.775	3	2
7 **Manufacture of wood products and furniture**	**.871**	**1.251**	**3**	**4**
14 Manufacture of wood products and furniture	.939	1.522	3	4
8 **Manufacture of paper and paperboard, and publishing**	**.632**	**.517**	**2**	**1**
15 Manufacture of paper and pulp, and publishing	.635	.519	2	2
9 **Manufacture of rubber, leather and products n.e.c.**	**.624**	**.807**	**2**	**2**
16 Manufacture of rubber products	.903	1.062	3	3
32 Manufacture of miscellaneous other products n.e.c.	.834	.731	3	2
10 **Manufacture of chemical and pharmaceutical products**	**.662**	**.613**	**2**	**2**
17 Manufacture of non-petrochemical chemicals	.883	.900	3	3
18 Manufacture of petrochemical products and petroleum	.741	.518	2	1
19 Manufacture of miscellaneous chemical products	.610	.786	2	3
20 Manufacture of pharmaceutical products and detergents	.294	.344	1	1
21 Manufacture of plastics products	.708	.691	2	2
11 **Manufacture of apparel and textiles**	**.621**	**.452**	**1**	**1**
22 Manufacture of textiles	.616	.650	2	2
23 Manufacture of apparel and apparel accessories	.539	.205	1	1
12 **Manufacture of footwear**	**3.051**	**2.562**	**5**	**5**
24 Manufacture of footwear and leather and fur products	2.306	2.386	4	4
13 **Manufacture of food, beverages, and ethyl alcohol**	**3.224**	**3.443**	**5**	**5**
25 Processing of coffee	3.481	2.833	5	5
26 Processing of plant products	3.326	3.496	5	5
27 Processing of meat, including slaughter	4.769	5.783	5	5
28 Processing of dairy products	.012	.045	1	1
29 Processing of sugar	4.309	10.085	5	5
30 Processing and refining of food fats and oils	12.427	10.151	5	5
31 Manufacture of other food products and beverages	2.062	1.852	4	4

Source: UN Comtrade 1990. Balassa (1965) comparative advantage is defined as $BADV_{i,t} \equiv (X^{\text{Brazil}}_{i,t} / \sum_k X^{\text{Brazil}}_{k,t})/(X^{\text{World}}_{i,t} / \sum_k X^{\text{World}}_{k,t})$ for sector i in year t, where $X_{i,t}$ are exports (5th quintile: strongest adv.).

Table 16: YEAR EFFECTS IN CONDITIONAL LOGIT ESTIMATION

	Separations			Accessions		
	(1)	(2)	(3)	(4)	(5)	(6)
Year 1990	-2.061	-2.126	-2.131	1.032	.963	.950
	(.136)***	(.145)***	(.145)***	(.126)***	(.131)***	(.131)***
Year 1991	-1.325	-1.356	-1.357	1.262	1.227	1.218
	(.067)***	(.070)***	(.070)***	(.062)***	(.064)***	(.064)***
Year 1992	-.970	-.980	-.979	1.101	1.089	1.084
	(.110)***	(.110)***	(.110)***	(.109)***	(.109)***	(.110)***
Year 1993	-.859	-.860	-.859	1.122	1.125	1.123
	(.067)***	(.067)***	(.067)***	(.067)***	(.067)***	(.067)***
Year 1994	-.863	-.858	-.858	.971	.983	.987
	(.047)***	(.047)***	(.047)***	(.047)***	(.047)***	(.047)***
Year 1995	-.445	-.432	-.433	.697	.720	.728
	(.085)***	(.086)***	(.086)***	(.086)***	(.087)***	(.087)***
Year 1996	-.378	-.368	-.368	.685	.699	.704
	(.050)***	(.050)***	(.050)***	(.052)***	(.052)***	(.052)***
Year 1997	-.204	-.194	-.194	.488	.501	.505
	(.039)***	(.040)***	(.040)***	(.041)***	(.041)***	(.041)***
Trade-related covariates						
2nd order interactions		yes			yes	
3rd order interactions		yes	yes		yes	yes

Source: *RAIS* 1990-98 (1% random sample), male workers nationwide, 25 to 64 years old, with manufacturing job. Separations and accessions exclude transfers, deaths, and retirements. Reference observations are employments with no reported separation or accession in a given year. Year effects from conditional logit estimation: column 1 completes column 6 of Table 9, columns 2 and 3 complete columns 2 and 3 of Table 14, column 4 completes column 6 of Table 10, columns 5 and 6 complete columns 5 and 6 of Table 14. Other regressors (not reported): Trade-related, sector (subsector IBGE level), plant and worker covariates. Robust standard errors in parentheses: * significance at ten, ** five, *** one percent.

Table 17: First-stage Predictions

	Separations			Accessions		
	Exp. Status	Prd. Mkt. Tariff	Imp. Pen.	Exp. Status	Prd. Mkt. Tariff	Imp. Pen.
	(1)	(2)	(3)	(4)	(5)	(6)
Instruments						
World imports APD	3.576 (.789)***	-2.278 (.097)***	-.011 (.053)	3.829 (.975)***	-2.121 (.111)***	.386 (.065)***
World imports CEE	43.712 (4.341)***	-33.870 (.534)***	-16.636 (.293)***	38.920 (5.551)***	-26.912 (.635)***	-17.067 (.370)***
World imports LAC	-4.740 (1.035)***	14.265 (.127)***	4.759 (.070)***	-2.022 (1.319)	14.041 (.151)***	4.865 (.088)***
World imports NAM	-2.380 (.525)***	-.652 (.065)***	-1.672 (.035)***	-2.468 (.662)***	.377 (.076)***	-1.992 (.044)***
World imports ODV	-2.142 (.763)***	-5.735 (.094)***	.312 (.052)***	-1.376 (.977)	-5.275 (.112)***	-.139 (.065)**
World imports OIN	4.173 (.957)***	-9.100 (.118)***	-5.678 (.065)***	3.977 (1.181)***	-10.354 (.135)***	-5.339 (.079)***
World imports WEU	13.940 (.461)***	2.158 (.057)***	1.953 (.031)***	14.437 (.564)***	1.469 (.065)***	2.095 (.038)***
USD Exch. Rate	.105 (.025)***	-.211 (.003)***	.011 (.002)***	.081 (.032)**	-.252 (.004)***	-.014 (.002)***
PPI Idx. EU	.703 (.115)***	-.928 (.014)***	.113 (.008)***	.974 (.144)***	-.941 (.016)***	.052 (.010)***
PPI Idx. NAM	.411 (.106)***	.850 (.013)***	-.120 (.007)***	.474 (.138)***	.802 (.016)***	-.200 (.009)***
Exogenous covariates						
Balassa Comp. Adv.	-.020 (.003)***	-.026 (.0003)***	-.022 (.0002)***	-.024 (.003)***	-.027 (.0004)***	-.022 (.0002)***
FDI Flow (USD billion)	.002 (.003)	.014 (.0004)***	.004 (.0002)***	.0002 (.004)	.014 (.0004)***	.005 (.0003)***
Herfindahl Index (sales)	.332 (.044)***	.048 (.005)***	.053 (.003)***	.252 (.054)***	-.026 (.006)***	.098 (.004)***
Log Employment	.052 (.002)***	.003 (.0002)***	-.0009 (.0001)***	.050 (.002)***	.003 (.0002)***	-.0007 (.0001)***
Share: Middle School or less	-.172 (.016)***	.008 (.002)***	-.007 (.001)***	-.184 (.017)***	.007 (.002)***	-.009 (.001)***
Share: Some High School	-.063 (.019)***	-.002 (.002)	.003 (.001)**	-.092 (.021)***	-.005 (.002)**	.002 (.001)
Share: White-collar occ.	.060 (.010)***	.006 (.001)***	-.002 (.0007)**	.057 (.012)***	.004 (.001)***	-.002 (.0008)**
F statistic (IV)	13.432	14,338.09	477.064	23.689	12,723.32	310.494

Sources: *WTF* (*NBER*) bilateral import data 1990-98; sector data 1990-98 from various sources at subsector IBGE level; *RAIS* 1990-98 labor force information; *SECEX* exporter information 1990-98. Weighted regressions using worker-sample observations (as in Table 9 for separations, Table 10 for accessions), controlling for year effects. Annual sector-weighted world imports, coefficients rescaled to imports in USD trillion. Robust standard errors in parentheses: * significance at ten, ** five, *** one percent.

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