Labor Reallocation in Response to Trade Reform*

Naércio Aquino Menezes-Filho
IBMEC São Paulo and Universidade de São Paulo

Marc-Andreas Muendler¶
University of California, San Diego and CESifo

November 20, 2007

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

---

*We thank Adriana Kugler, Steve Redding, T.N. Srinivasan, Lars Vilhuber, Nina Pavcnik and Lori Kletzer for insightful discussions, and seminar and conference participants at Universidad de los Andes, Purdue, Stanford, IPEA Rio de Janeiro, Universidade de Brasília, the Munich-Tübingen Trade Workshop, ANPEC, LACEA Mexico, NBER-CAFE Nuremberg, the Econometric Society summer meetings Durham, the NBER summer institute, UC Santa Cruz, UNC-Chapel Hill and Vanderbilt for their comments. We thank Mary Amiti and Luis Servén for trade data, and Paulo Furtado and the Brazilian Ministry of Labor for assistance with *RAIS*. We thank Alexandre Brandão and Aline Visconti at IBGE for tabulations of *PIA*. Jennifer Poole and Andrea Curi provided superb research assistance. Muendler acknowledges NSF support (SES-0550699) with gratitude. In-depth statistics beyond this paper are available from URL *econ.ucsd.edu/muendler/research*.

¶muendler@ucsd.edu (*www.econ.ucsd.edu/muendler*). Ph: +1 (858) 534-4799.

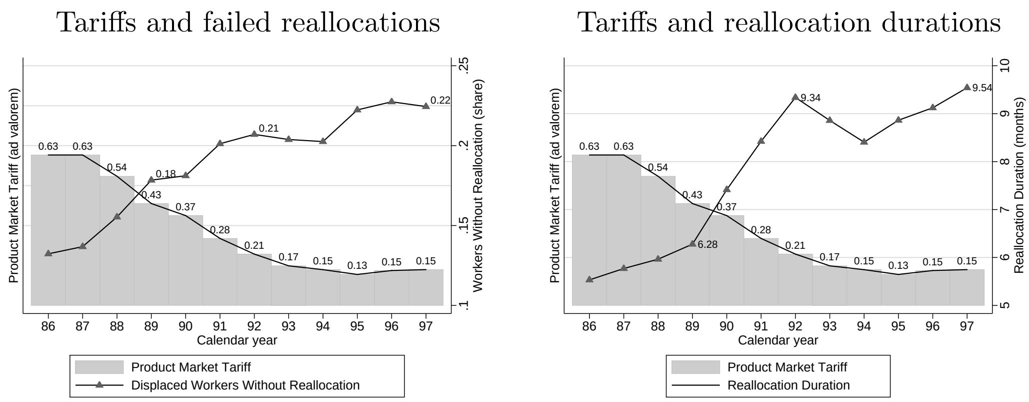
1 Introduction

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 (Hay 2001, Pavcnik 2002) 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.

The 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, where idle resources result. 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

[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). Santos-Paulino and Thirlwall (2004) point to the paucity of research into mechanisms by which trade reforms affect economic performance.



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 (2003), employment weighted at *Nível 50* sector level.

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

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

When there are frequent transitions between formal employment and other types of work, it takes linked employer-employee data to measure reallocation directions and durations after trade reform and to assess the overall economic performance. 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 suggest that tariff reductions are related to more transitions out of formal 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 forthcoming). 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

[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.

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.

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 is organized as follows. 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 analyzes worker separations and accessions to identify sector and firm predictors that explain reallocation delays and failures. Section 5 subjects these predictions to numerous robustness checks, including work status transitions from a household perspective. Section 6 discusses implications for trade theory. Section 7 concludes.

2 Data

We track Brazil's labor reallocation with two main data sets. From both data sets, 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.[5] 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 and track the workers across their identified employers over time so that we can cover national formal-sector migration and estimate worker-fixed effects.[6] To *RAIS*, we

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

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

[6] 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.

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.

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.[7] Overall, formal-sector employment increases from 23.2 million to 24.5 million between 1990 and 1998 in *RAIS*.

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 prime-age male workers' IDs, and then trace the selected workers through their formal jobs. For most separation statistics, we remove multiple jobs and only retain a worker's highest paying job at a given moment.

Our concern is with potentially idle labor: displaced workers who await formal-sector reallocation. They are not directly observable in *RAIS*. However, *RAIS* records two margins that change the pool of prime-age male workers to be reallocated: separations from formal jobs fill the pool, and accessions into formal jobs empty the pool of workers to be reallocated. When we infer separations and

[7] 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.

Table 1: Labor Market Performance and Economic Outcomes

	1986	1990	1992	1994	1998
FAILED REALLOCATIONS WITHIN A YEAR					
Mean failure rate (share of displaced)	.248	.323	.410	.369	.459
young workers	.235	.303	.354	.326	.366
college-educated workers	.258	.315	.350	.337	.387
Change over 1990		.000	.086	.046	.136
Idle labor (foregone share of GDP)		.000	.014	.006	.024
DURATIONS OF SUCCESSFUL REALLOCATIONS WITHIN A YEAR					
Mean duration (in months)	2.776	3.808	4.206	4.108	4.220
young workers	2.226	3.135	3.460	3.262	3.367
college-educated workers	1.691	2.429	2.423	2.250	2.282
Change over 1990 (one twelfth)		.000	.033	.025	.034
Idle labor (foregone share of GDP)		.000	.005	.003	.006

Sources: *RAIS* 1986-1999 (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 12 months (*upper panel*) or rehired into a formal-sector job within 12 months (*lower panel*). *PME* 1986-1999, share of idle workers (unemployed or withdrawn from labor force), and *Banco Central do Brasil*, GDP. We define young workers to have ten or less years of potential labor force experience, and college-educated workers to have some college education. Foregone GDP is the unrealized wage bill, measured as the product of the observed change over 1990 times the number of newly displaced workers during the year times their wage upon displacement. Idle labor is defined as the share of displaced workers with transitions to unemployment or out of the labor force.

accessions, we exclude transfers across plants within the same firm, as well as retirements and reported deaths on the job. An accession is defined as a worker's hiring into the first formal employment in the calendar year. Conversely, we define a separation as a worker's quit or layoff from the last formal employment in the calendar year. Among the separations, quits are infrequent compared to layoffs (Table 2). We consider separations as a single category for regression analysis, where we detect no marked difference between quits and layoffs for trade-related predictors.

Two important measures for the potential idleness of labor are the durations of successful reallocations within a given time period, such as twelve months following displacement, and the rate of failed reallocations within the time horizon. Table 1 documents that the share of displaced workers without reallocation for a year almost doubles from 25 to 46 percent between 1986 and 1998. There is some variation in the failure rate across skill groups within any given year:

young and college-educated workers' reallocations fail less frequently than average. Time variation, however, dwarfs the skill-group differences. A similar pattern applies to durations of successful reallocations in the lower panel of Table 1. The relatively minor cross-sectional differences between skill groups, compared to major time variation, suggests that studying macroeconomic sources of variation in labor-market performance promises to uncover first-order changes in labor-market outcomes.

Idle resources in the labor market are a foremost component of Brazil's aggregate performance. For reallocation failures in the upper panel of Table 1, 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 the share of displaced formal-sector workers as idle who typically become unemployed or move out of the labor force—a 36 percent share on average in *PME* 1990-98. 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 in the lower panel of Table 1 implies another half percent of foregone GDP in 1992 and .6 percent in 1998. This brings the total foregone wage bill to almost 2 percent of GDP in 1992, to almost 1 percent in 1994 (a year with strong GDP growth), and to 3 percent in 1998. The estimates are conservative because we only consider two out of five employees, prime-age male workers, and we assume that displaced workers who become informal or self-employed retain the full pre-displacement earnings immediately after displacement.

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), surveyed under a rotating panel similar to the U.S. *PSID* and the British *BHPS*. We use 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 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. 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 sectors exhibit an average increase in informality of only around fifty percent.

Sector data. 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.[8] We combine sector-level variables from several sources with *RAIS* and *PME*. We obtain *ad valorem* tariffs by sector and year and use input-output matrices to compute intermediate-input tariffs in addition to product-market tariffs (see Appendix). Figure 2 shows that the steep tariff schedule of 1990 is replaced with a flatter schedule by 1997. The initially most protected sectors are subject to the strongest declines in tariffs.

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

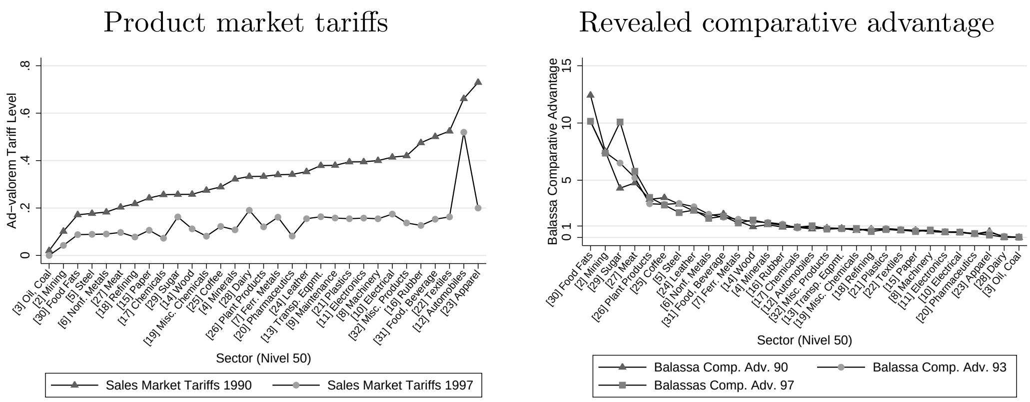
$$BADV_{i,t} \equiv \frac{X_{i,t}^{\text{Brazil}} / \sum_k X_{k,t}^{\text{Brazil}}}{X_{i,t}^{\text{World}} / \sum_k X_{k,t}^{\text{World}}},$$

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. The right panel in Figure 2 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.[9] Figure 2 illustrates the regression re-

[8] 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 at URL *econ.ucsd.edu/muendler/research*. Results exhibit little sensitivity to alternative regional samples or sector classifications.

[9] Results are at URL *econ.ucsd.edu/muendler/research*.



Source: Product tariffs from Kume et al. (2003) and UN Comtrade 1986-98. Left panel: Sectors at *Nível 50* ordered by 1990 product tariff. Right panel: Sectors at *Nível 50* ranked by sector-fixed Balassa comparative advantage; 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 2: **Manufacturing Tariffs and Balassa Comparative Advantage**

sults. 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.[10] Additional instruments are components of the sectoral real exchange rate: the U.S. dollar exchange rate and sector price levels in the United States

[10] 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).

and the EU. We denote exchange rates as foreign divided by domestic prices so that a high exchange rate means a depreciated level. These instruments prove to be significant covariates of trade-reform variables but, being largely unpredictable themselves to workers and firms, the instruments are plausibly unrelated to job matches.

Firm data. We combine the linked employer-employee data from *RAIS* with additional firm-level data.[11] 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.[12] 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* (see Appendix). Confidentiality requirements only allow us to use randomly combined three-firm cells from *PIA*.

Workforce characteristics and trade exposure. Table 2 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 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.[13] 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, re-

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

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

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

Table 2: *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 Productivity	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.

spectively, 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 over-represented. 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 for brevity), there are remarkable mean differences between an exporter and an average firm.[14] 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 productiv-

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

ity in 1998 exceeds log labor productivity in 1990 by 4.5 percent in the estimation sample, and by 4.7 percent at manufacturing exporters.

3 Labor Reallocation, Productivity 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. The evidence documents chief mechanisms of resource reallocation after Brazil's trade reform.

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 3 points to the key explanation for Brazil's labor-market experience after trade reform. The table 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 relative log TFP change of 3.5 percent between 1990 and 1998 is modest (column 1).[15] Substantial capital accumulation contributes to the faster increase in log labor productivity by 7.3 percent between 1990 and 1998 (column 5). Alongside, Table 3 reports the raw covariance of year-

[15] In Table 3, 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.

Table 3: 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).

over-year productivity changes at surviving firms (columns 4 and 8)—a term in the Haltiwanger (1997) decomposition over time.[16]

The decompositions in Table 3 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).[17] These facts are well known for Brazil and similar countries, but sometimes 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).[18] So, firms with increasing productivity expand output shares but reduce employment. Resource

[16]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.

[17]Winters (2004) surveys the broader literature on trade liberalization and economic growth and concludes that trade liberalization typically raises growth, mostly through productivity change.

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

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 a substantial 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 2003). By contrast, far-reaching trade reform under the Collor administration in 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. We control for effective protection and use both product-market and intermediate-input tariffs in our subsequent analysis.

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.

Table 4: 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
Stationary distribution	3.4	3.1	4.0	9.7	4.9	52.9	21.9	100.0
Adjusted Stationary distrib.	2.5	2.3	2.9	7.0	3.6	38.2	43.5	100.0

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

Sources: *RAIS* 1986, 1990, 1994 and 1998 (1% random sample), male workers nationwide, 25 to 64 years old; and *PME* 1986-1999. 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. The stationary distribution is the normalized left eigenvector of the *RAIS* transition matrix associated with the eigenvalue of one; the failure adjusted stationary distribution is the eigenvector based on an estimate of 4-year failure-to-failure transitions from *PME* (63.6% of non-formal *PME* workers are in non-formal work status after three annual transitions, replacing the zero from *RAIS*).

Worker reallocation. Table 4 reports transitions of displaced prime-age male workers from formal-sector jobs to other formal-sector jobs at the four-year 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 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 the table.

The dominant fraction of workers with displacement from a traded-goods in-

[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.

dustry, 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 four years (column 7). Repeating the statistical exercise for various subperiods shows that reallocation patterns in traded-goods industries are remarkably stable. 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.[20]

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 far exceed those off the diagonal (column 1-5). These within-sector reallocation rates also far exceed the stationary distribution of employment in the sector (reported in the last rows of the Table), rendering within-sector reallocation transitory. Repeating the exercise at the annual horizon shows that diagonal entries are around seven percent larger at the four-year horizon than within a year. This suggests that successful reallocations over the medium to long term are within-sector shifts, in contrast to classic trade theory which posits that the largest fraction of reallocated workers should move to the high-quintile industries (columns 4 and 5) from every separation sector.

Recent trade models emphasize the importance of intra-industry heterogeneity across firms.[21] Our linked employer-employee data allow us to track worker reallocations across identified firms. During the 1990s, around a million prime-age male workers are successfully reallocated every year. See Table 5. Nine in ten workers shift to nonexporters after separation, while one in ten transitions to an exporter. The relatively small magnitude of transitions to exporters, with less than one in ten displaced workers moving to an exporter overall, and an observed rehiring bias at exporters towards former exporter workers, suggest that labor shifts from nonexporters to exporters are not a major channel of worker reallocation.[22]

---

[20]In comparison, only 39.7 percent of *PME* household members in 1998 have a formal job; the remaining 61.3 percent hold an informal job (15.3), are self-employed (23.7), unemployed (5.0) or out of the labor force (16.4 percent). The estimate of the adjusted stationary distribution in the last row of Table 4 suggests that 43.5 percent of the once-formal workers become non-formal in steady state, and contribute to the overall pool of more than 60 percent of household members who work under a non-formal status.

[21]See Greenaway and Kneller (2007) for a survey.

[22]Reallocations within firms are minor (Muendler forthcoming): 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

Table 5: 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.

The descriptive evidence so far is based on unconditional means. The remainder of the paper subjects reallocation statistics to multivariate controls and instrumental-variable estimation.

4 Formal-Sector Separations and Accessions

To understand determinants of labor reallocation in the formal sector, we turn to industry, plant, job and worker characteristics as explanatory variables for separations and accessions. Employers adjust workforces through worker separations and accessions. Separations in turn burden, and accessions unburden, the pool of workers to be reallocated.

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:

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

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 $S(J(i))$, including a sector-fixed effect in some specifications; $\mathbf{y}_{J(i),t}$ 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

between plants within their employing firm.

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. For theoretical consistency with random shocks to employer-employee matches, the disturbance is assumed to be logistic and independent across matches. We fit this conditional logit model (1) 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 6 presents conditional logit estimates of separations from formal manufacturing jobs, where the conditioning removes worker-fixed effects (worker-FE logit). 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 6). FDI inflows into the sector predict a statistically significant reduction in displacement 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 5), 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

Table 6: WORKER-FIXED EFFECT LOGIT ESTIMATION OF SEPARATIONS

	(1)	(2)	(3)	(4)	(5)	(6)
Balassa Comp. Adv.	.080			.169	.204	-.094
	(.021)***			(.024)***	(.023)***	(.049)*
Exporter Status		.289		.283	.301	.284
		(.028)***		(.028)***	(.028)***	(.028)***
Product Market Tariff			-.104	-.705	-1.383	-2.361
			(.416)	(.426)*	(.410)***	(.476)***
Intm. Input Tariff			1.601	2.880	-1.420	5.149
			(.633)**	(.678)***	(.553)**	(.748)***
Import Penetration				1.257	6.035	3.227
				(.388)***	(.349)***	(.638)***
Sector-level covariates						
Sector real exch. rate	.733	.843	.353	-.398	.213	-1.224
	(.624)	(.626)	(.640)	(.645)	(.069)***	(.699)*
FDI Flow (USD billion)	-.025	-.012	-.018	-.048	.047	-.039
	(.020)	(.020)	(.020)	(.020)**	(.019)**	(.020)**
Herfindahl Index (sales)	-.371	-.517	-.399	-.354	.929	.881
	(.317)	(.316)	(.329)	(.343)	(.320)***	(.639)
Plant-level covariates						
Log Employment	-.343	-.370	-.341	-.377	-.410	-.383
	(.011)***	(.011)***	(.011)***	(.011)***	(.011)***	(.011)***
Share: Middle School or less	-.750	-.658	-.719	-.663	-.793	-.692
	(.131)***	(.131)***	(.131)***	(.132)***	(.129)***	(.132)***
Share: Some High School	-.444	-.392	-.440	-.393	-.214	-.413
	(.148)***	(.148)***	(.147)***	(.148)***	(.145)	(.148)***
Share: White-collar occ.	.721	.700	.739	.691	.552	.683
	(.075)***	(.074)***	(.074)***	(.075)***	(.073)***	(.075)***
Worker-level covariates						
Tenure at plant (in years)	1.367	1.350	1.362	1.351	1.390	1.351
	(.036)***	(.036)***	(.036)***	(.036)***	(.037)***	(.036)***
Pot. labor force experience	.006	.006	.006	.006	.031	.006
	(.002)**	(.002)**	(.002)**	(.002)**	(.002)***	(.002)**
Unskilled Wh. Collar Occ.	-.256	-.251	-.259	-.262	-.199	-.267
	(.067)***	(.067)***	(.067)***	(.067)***	(.065)***	(.067)***
Year effects	yes	yes	yes	yes		yes
Sector effects						yes
Obs.	145,408	145,408	145,408	145,408	145,408	145,408
Pseudo R^2	.148	.149	.148	.150	.137	.151

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.

in important ways.

Inclusion of sector-fixed effects removes unobserved sectoral differences that potentially co-determine separations (column 6). The sector effects control for potential differences in the effect of labor institutions, for instance, whose reform in 1988 precedes trade liberalization in 1990. Expectedly, inclusion of sector indicators turns the coefficient on comparative advantage, which is highly sector specific and largely time invariant, insignificant. For the other trade regressors, however, coefficient estimates increase in absolute value (compared to column 4) and remain highly significant. We subsequently emphasize the more conservative estimates without sector effects.

Before discussing plant and worker-level variables, we turn to the opposite margin: Table 7 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 4). 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 13). 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 5), 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. Lower input tariffs, which tend to make competition less fierce, predict more accessions. Lower output tariffs and the arrival of additional imports, which tend to make competition more fierce, are associated with fewer accessions. When we condition on both year and sector effects (column 6), comparative advantage does expectedly not turn significant, whereas coefficients for all other trade regressors increase in absolute value (compared to column 4) and remain or become highly significant. As we do for separations, we therefore base much of our subsequent discussion on the more conservative estimates without sector effects.

Larger manufacturing plants exhibit less turnover: they displace significantly

Table 7: Worker-fixed Effect Logit Estimation of Accessions

	(1)	(2)	(3)	(4)	(5)	(6)
Balassa Comp. Adv.	.041			-.016	-.114	-.067
	(.017)**			(.020)	(.019)***	(.048)
Exporter Status		-.449		-.439	-.429	-.438
		(.027)***		(.027)***	(.026)***	(.027)***
Product Market Tariff			1.306	1.246	2.474	1.822
			(.379)***	(.393)***	(.379)***	(.498)***
Intm. Input Tariff			-3.258	-3.073	-3.846	-2.954
			(.540)***	(.598)***	(.514)***	(.750)***
Import Penetration				.198	-3.919	1.764
				(.355)	(.307)***	(.665)***
Sector-level covariates						
Sector real exch. rate	-1.264	-.955	-.953	-.810	.038	-.844
	(.605)**	(.606)	(.626)	(.639)	(.076)	(.718)
FDI Flow (USD billion)	.039	.047	.056	.058	.031	.058
	(.022)*	(.021)**	(.021)***	(.022)***	(.021)	(.022)***
Herfindahl Index (sales)	-.348	-.344	-.795	-.788	-2.335	-.838
	(.268)	(.268)	(.282)***	(.297)***	(.277)***	(.655)
Plant-level covariates						
Log Employment	-.190	-.140	-.189	-.141	-.112	-.138
	(.008)***	(.009)***	(.008)***	(.009)***	(.008)***	(.009)***
Share: Middle School or less	.947	.857	.940	.850	.828	.849
	(.107)***	(.105)***	(.107)***	(.105)***	(.104)***	(.105)***
Share: Some High School	.740	.667	.739	.668	.468	.668
	(.124)***	(.122)***	(.124)***	(.122)***	(.120)***	(.122)***
Share: White-collar occ.	-.675	-.614	-.679	-.621	-.534	-.625
	(.067)***	(.067)***	(.067)***	(.067)***	(.064)***	(.067)***
Worker-level covariates						
Prof. or Manag'l. Occ.	-.801	-.807	-.801	-.807	-.827	-.810
	(.068)***	(.068)***	(.068)***	(.068)***	(.066)***	(.068)***
Tech'l. or Superv. Occ.	-.603	-.610	-.597	-.604	-.623	-.601
	(.064)***	(.064)***	(.064)***	(.064)***	(.062)***	(.064)***
Unskilled Wh. Collar Occ.	-.490	-.497	-.488	-.495	-.519	-.497
	(.061)***	(.062)***	(.062)***	(.062)***	(.060)***	(.062)***
Skilled Bl. Collar Occ.	-.417	-.413	-.413	-.410	-.443	-.410
	(.032)***	(.032)***	(.032)***	(.032)***	(.031)***	(.032)***
Year effects	yes	yes	yes	yes		yes
Sector effects						yes
Obs.	112,974	112,974	112,974	112,974	112,974	112,974
Pseudo R^2	.036	.040	.037	.041	.026	.042

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.

Table 8: Trade Exposure and Predicted Labor Market Outcomes

	1990	1992	1994	1998
Trade Exposure				
Import Penetration	.041	.056	.060	.103
Product Market Tariff	.358	.202	.142	.167
Intm. Input Tariff	.278	.152	.107	.129
Change in Separation rates predicted by				
change in Import Penetration since 1990		.016	.020	.064
changes in Tariffs since 1990		.067	.092	.081
Change in Accession rates predicted by				
change in Import Penetration since 1990		-.008	-.012	-.040
changes in Tariffs since 1990		.018	.023	.019

Source: *RAIS* 1990-98, male workers nationwide, 25 to 64 years old, with manufacturing job (estimation samples from Tables 6 and 7). Sector information at subsector IBGE level. Predicted changes in separation and accession rates based on marginal effects implied by column (6) estimates in Tables 6 and 7 ($\hat{P}(1-\hat{P})$ is .170 for separations and .174 for accessions).

fewer (Table 6) and they hire significantly fewer workers (Table 7). Plants with less educated workforces and more blue-collar jobs separate from workers significantly 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.

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 6 and 7 (columns 5). Import penetration more than doubles between 1990 and 1998, while product-market and input tariffs drop by more than half. As Table 8 reports, 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 mean separation rate of .282 over the 1990s (Table 2). The manufacturing sector employs roughly one in four male workers in the national labor force, so these predictions 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.

5 Concomitant Economic Changes and Reforms

The evidence so far shows that Brazil's trade reform predicts salient changes to worker separations and accessions. 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.[23] This section addresses 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.

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

[23] 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).

Table 9: LINEAR AND INSTRUMENTAL-VARIABLE WORKER-FE 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 4 in Tables 6 and 7. Further regressors (not reported): Year indicators, sector, plant and worker covariates. Robust standard errors in parentheses: * significance at ten, ** five, *** one percent.

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, calculate the destination-region imports from other source-countries than Brazil, and weight the destinations' importance for Brazil's industries using Brazilian exports in 1990. 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 current foreign producer price levels in industrialized economies are arguably unrelated to the termination and formation of job matches in 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:[24]

$$\sigma_{i,t} = \hat{\mathbf{z}}_{S(J(i)),t}\beta_z + \hat{\mathbf{y}}_{J(i),t}\beta_y + \mathbf{x}_{i,t}\beta_x + \alpha_i + \alpha_t + \epsilon_{i,t}, \tag{2}$$

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 (1). 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 (2). 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.

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 14 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.[25] 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 9 redisplays conditional logit estimates for separations and accessions in columns 1 and 4 (from Tables 6 and 7, column 4). 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

[24]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).

[25]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.

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.[26] 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 investment inflows and policy shifts such as macroeconomic stabilization, capital-account liberalization, and privatization.[27] 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 sector 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 (1).

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 (1) 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 10 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

[26]In instrumental-variable regressions with sector-fixed effects, more trade-related predictors lose significance but there is no sign reversal.

[27]Labor-market institutions were altered preceding trade reform and their industry-specific impact is controlled for with sector-fixed effects (Tables 6 and 7, column 6).

Table 10: Alternative Worker-fixed Effect Logit Specifications

	Cdl. logit	Young worker	Primary school	College educ.	Privat. control	Outsrc. job ind.
	(1)	(2)	(3)	(4)	(5)	(6)
			Separations			
Balassa Comp. Adv.	.169	.498	.145	.216	.170	.169
	(.024)***	(.267)*	(.028)***	(.150)	(.026)***	(.024)***
Exporter Status	.283	.379	.296	.297	.283	.283
	(.028)***	(.243)	(.033)***	(.143)**	(.028)***	(.029)***
Product Market Tariff	-.705	-3.960	-.500	-1.771	-.694	-.751
	(.426)*	(4.290)	(.499)	(2.281)	(.427)	(.430)*
Intm. Input Tariff	2.880	10.027	2.469	7.146	2.875	3.010
	(.678)***	(7.163)	(.779)***	(4.086)*	(.675)***	(.686)***
Import Penetration	1.257	8.588	.678	.886	1.264	1.269
	(.388)***	(3.668)**	(.477)	(1.995)	(.392)***	(.391)***
addl. regressor(s)					-.142	-.018
					(1.227)	(.037)
Obs.	145,408	2,897	110,831	7,498	145,408	143,536
Pseudo R^2	.150	.391	.161	.245	.150	.151
			Accessions			
Balassa Comp. Adv.	-.016	-.120	-.006	-.141	-.024	-.015
	(.020)	(.209)	(.023)	(.118)	(.022)	(.021)
Exporter Status	-.439	-.477	-.420	-.776	-.439	-.437
	(.027)***	(.216)**	(.031)***	(.140)***	(.027)***	(.027)***
Product Market Tariff	1.246	.099	1.333	2.033	1.118	1.185
	(.393)***	(3.290)	(.451)***	(2.092)	(.412)***	(.397)***
Intm. Input Tariff	-3.073	-7.113	-2.943	-5.152	-2.987	-3.041
	(.598)***	(5.668)	(.673)***	(3.393)	(.603)***	(.604)***
Import Penetration	.198	-9.315	.084	-.720	.128	.181
	(.355)	(3.845)**	(.423)	(1.948)	(.363)	(.358)
addl. regressor(s)					1.140	-.098
					(1.166)	(.033)***
Obs.	112,974	2,752	86,468	4,786	112,974	110,985
Pseudo R^2	.041	.223	.043	.088	.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 4 in Tables 6 and 7. Further regressors (not reported): Year indicators, sector, plant and worker covariates. Robust standard errors in parentheses: * significance at ten, ** five, *** one percent.

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.

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 attribution. 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 (1). As column 5 in Table 10 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 (1) results in no statistically significant coefficient change (column 6). 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.

The constitutional labor-market reforms in 1988 precede trade liberalization in 1990. The strengthened results in regressions with sector-fixed effects, and the unaltered evidence from instrumental-variable regressions, render it little plausible that changes to labor institutions can be erroneously attributed to trade.

Firm-level labor productivity. Exporters are more productive than nonexporters (Table 2). 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 (1). For this purpose, we use the

Table 11: WORKER-FE 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)***
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. Estimates in column 1 and 4 repeat column 4 in Tables 6 and 7. Further regressors (not reported): Year indicators, sector, plant and worker covariates. Robust standard errors in parentheses: * significance at ten, ** five, *** one percent.

subsample of *RAIS* firms that are surveyed in *PIA*, for which firm-level labor productivity is inferrable. This link reduces the number of observations markedly. Moreover, confidentiality requirements only allow us to use randomly combined three-firm cells from *PIA*, resulting in a loss of efficiency.

Table 9 redisplays conditional logit estimates for separations and accessions on the full sample in columns 1 and 4 (from Tables 6 and 7, column 4). 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 3) that faster labor productivity growth at manufacturing firms correlates 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 clearly reinforces several findings.

Work status transitions. Household data allow us to discern in more detail the categories of work-status transitions out of formal employment. We estimate a multinomial logit model of a single work status transition for every *PME* household member at the annual horizon.[28] 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; (4) the worker moves to unemployment; and (5) the worker withdraws from the labor force.

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

$$Pr(\sigma_{i,t+1} |\, \sigma_{i,t} = \sigma; \mathbf{x}, \mathbf{z}) = \frac{\exp\{\mathbf{z}_{S(i),t}\beta_z^\sigma + \mathbf{x}_{i,t}\beta_x^\sigma + \alpha_t^\sigma + \alpha_{c(i),t}^\sigma\}}{\sum_{\varsigma\in\mathbb{S}} \exp\{\mathbf{z}_{S(i),t}\beta_z^\varsigma + \mathbf{x}_{i,t}\beta_x^\varsigma + \alpha_t^\varsigma + \alpha_{c(i),t}^\varsigma\}}, \qquad (3)$$

where $\mathbf{z}_{S(i),t}$ is a vector of sector-level covariates of the household member's initial sector $S(i)$, including a sector-fixed effect in some specifications; $\mathbf{x}_{it}$ 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 and independent across employer-employee matches. As a rudimentary version of a worker-fixed effect, we include among the job-worker covariates 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 (3) 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 $t+1$.

Table 12 presents predictions without sector-fixed effects in the upper panel,

[28] 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 intrinsic ordering.

Table 12: 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)
	No sector-fixed effects			
Product Market Tariff	-2.842	-4.016	-2.080	-.129
	(.799)***	(.803)***	(.906)**	(.850)
Intm. Input Tariff	1.823	4.250	1.849	1.089
	(.974)*	(.973)***	(1.102)*	(1.037)
Formal empl. for four months	-1.679	-1.307	-.736	-1.032
	(.072)***	(.078)***	(.103)***	(.090)***
Some High School	-.349	-.370	-.271	.217
	(.079)***	(.077)***	(.085)***	(.087)**
Some College	-.464	-.639	-.651	.449
	(.205)**	(.228)***	(.231)***	(.232)*
College Degree	-.724	-.520	-1.096	-.172
	(.146)***	(.140)***	(.183)***	(.181)
Obs.	25,520			
Pseudo R^2	.06			
	Sector-fixed effects			
Product Market Tariff	-.319	-1.387	-2.019	-1.960
	(1.463)	(1.466)	(1.653)	(1.569)
Intm. Input Tariff	-.187	2.699	1.538	2.407
	(1.540)	(1.538)*	(1.720)	(1.626)
Formal empl. for four months	-1.626	-1.282	-.737	-1.051
	(.072)***	(.079)***	(.104)***	(.091)***
Some High School	-.299	-.361	-.266	.224
	(.079)***	(.078)***	(.086)***	(.088)**
Some College	-.399	-.628	-.626	.472
	(.206)*	(.229)***	(.233)***	(.233)**
College Degree	-.635	-.509	-1.086	-.157
	(.148)***	(.141)***	(.185)***	(.182)
Obs.	25,520			
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. Sector-level variables at subsector IBGE level. Controlling for year and city effects as well as a worker's potential labor-market experience in both panels, for sector effects in lower panel. Robust standard errors in parentheses: * significance at ten, ** five, *** one percent.

and with sector-fixed effects in the lower panel.[29] In both panels of Table 12, workers with stable formal-sector employment for four months are significantly less likely to lose formality status over the following year. Higher educational attainment, from some high-school attainment through college education, predicts significantly fewer transitions into informality, self employment and unemployment. Education groups with less than a college degree are more likely to transition out of the labor force.

In the upper panel, lower product-market tariffs are associated with significantly higher odds of transitions into informality, self employment, and unemployment. Intermediate-input tariff coefficients show converse signs, and predict significantly more transitions into self employment. 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 alleviate competitive pressure. Overall, the estimates predict that trade liberalization results in significantly more transitions out of formality. When sector-fixed effects at the subsector IBGE level are included among the regressors in the lower panel, the product-market tariff loses statistical significance but exhibits no sign reversal, and similarly for most intermediate-input tariff coefficients.

In related research, Goldberg and Pavcnik (2003) detect no significant effect of trade liberalization on the incidence of informality in sector data for Brazil. Estimates in the lower panel of Table 12, too, lack statistical significance after inclusion of sector-fixed effects. The worker-level evidence from *PME* nevertheless remains consistent with the hypothesis that sector differences in tariffs are associated with changes in transitions out of formality. The coefficient signs, and their statistical significance in the upper panel where tests on sector variables are more powerful, agree with our detailed evidence from *RAIS*. *RAIS* allows us to control for worker-fixed effects over long horizons as well as for employer characteristics, in addition to sector-fixed effects, and we find significantly more formal-sector separations and significantly fewer formal-sector accessions after trade reform. Given overall formal-sector employment growth between 1990 and 1998, the net effect must be an increase in non-formal work status.

[29] For consistency with our evidence from *RAIS*, we map the *PME* sector definitions to the subsector IBGE level. Results are similar under alternative sector definitions (see our working paper Menezes-Filho and Muendler 2007).

6 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.[30] 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 Brazil's mean trading partner shows that Brazil's labor market is considerably more rigid.[31] So, those theories would predict Brazil to specialize in

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

[31] For the World Bank's four rigidity and difficulty indices (hiring difficulty, hours rigidity, firing difficulty, employment rigidity) and its firing-cost measure, 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 at URL *econ.ucsd.edu/muendler/research*.

Table 13: Worker-fixed Effect 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. Columns 1 and 4 repeat column 4 of Tables 6 and 7. Further regressors (not reported): Year indicators, sector, plant and worker covariates. Robust standard errors in parentheses: * significance at ten, ** five, *** one percent.

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 firms.[32]

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 13 compares our previous separation and accessions estimates in columns 1 and 4 (from Tables 6 and 7, column 6) to regressions with interaction terms in the remaining columns. There are no remarkable changes to coefficient estimates for separations. At the 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. As Arbache, Dickerson and Green (2004) have argued in the context of relative wage responses to trade before, standard trade theory ignores trade-induced technology adoption and implied relative labor-demand changes. Once augmented to accommodate productivity change, Heckscher-Ohlin-Samuelson textbook models 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 theoretical research provides novel 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,

[32]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.

and provides stronger managerial incentives to raise production efficiency. If factor productivity rises faster than output in a general-equilibrium extension to the Raith (2003) model, 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.

7 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 workers across employers and industries in the aftermath of Brazil's large-scale trade reform. The data reveal 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. 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 out of formality, 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. *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. 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 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 (undocumented) 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 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.[33] We map ISCO-88 categories to *RAIS* occupations as follows:

[33]See documentation at URL *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 Metropolitan household data

The Brazilian monthly employment survey *PME* (*Pesquisa Mensal de Emprego*) is conducted by Brazil's statistical bureau IBGE, using a rotating panel. *PME* follows households for 16 months, with an eight-month interval after the fourth interview.[34] 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. 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.

C Manufacturing firm data

For descriptive statistics in Table 3 and robustness checks in Table 11, 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

[34]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 $m+15$.

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*.

D Additional sector data

We use data on *ad valorem* tariffs by sector and year from Kume, Piani and Souza (2003). The tariffs are the legally stipulated nominal rates for Brazil's trade partners with no preferential trade agreement, and not weighted by source country. 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 intermediate-input tariff as the weighted arithmetic average of the product-market tariffs, using sector-specific shares of inputs for the input-output matrix as weights.

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 output less net exports, the relevant domestic environment in which they compete. We define the effective rate of market penetration as imports per absorption. 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 14: 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	-2.278	-.011	3.829	-2.121	.386
	(.789)***	(.097)***	(.053)	(.975)***	(.111)***	(.065)***
World imports CEE	43.712	-33.870	-16.636	38.920	-26.912	-17.067
	(4.341)***	(.534)***	(.293)***	(5.551)***	(.635)***	(.370)***
World imports LAC	-4.740	14.265	4.759	-2.022	14.041	4.865
	(1.035)***	(.127)***	(.070)***	(1.319)	(.151)***	(.088)***
World imports NAM	-2.380	-.652	-1.672	-2.468	.377	-1.992
	(.525)***	(.065)***	(.035)***	(.662)***	(.076)***	(.044)***
World imports ODV	-2.142	-5.735	.312	-1.376	-5.275	-.139
	(.763)***	(.094)***	(.052)***	(.977)	(.112)***	(.065)**
World imports OIN	4.173	-9.100	-5.678	3.977	-10.354	-5.339
	(.957)***	(.118)***	(.065)***	(1.181)***	(.135)***	(.079)***
World imports WEU	13.940	2.158	1.953	14.437	1.469	2.095
	(.461)***	(.057)***	(.031)***	(.564)***	(.065)***	(.038)***
USD Exch. Rate	.105	-.211	.011	.081	-.252	-.014
	(.025)***	(.003)***	(.002)***	(.032)**	(.004)***	(.002)***
PPI Idx. EU	.703	-.928	.113	.974	-.941	.052
	(.115)***	(.014)***	(.008)***	(.144)***	(.016)***	(.010)***
PPI Idx. NAM	.411	.850	-.120	.474	.802	-.200
	(.106)***	(.013)***	(.007)***	(.138)***	(.016)***	(.009)***
Exogenous covariates						
Balassa Comp. Adv.	-.020	-.026	-.022	-.024	-.027	-.022
	(.003)***	(.0003)***	(.0002)***	(.003)***	(.0004)***	(.0002)***
FDI Flow (USD billion)	.002	.014	.004	.0002	.014	.005
	(.003)	(.0004)***	(.0002)***	(.004)	(.0004)***	(.0003)***
Herfindahl Index (sales)	.332	.048	.053	.252	-.026	.098
	(.044)***	(.005)***	(.003)***	(.054)***	(.006)***	(.004)***
Log Employment	.052	.003	-.0009	.050	.003	-.0007
	(.002)***	(.0002)***	(.0001)***	(.002)***	(.0002)***	(.0001)***
Share: Middle School or less	-.172	.008	-.007	-.184	.007	-.009
	(.016)***	(.002)***	(.001)***	(.017)***	(.002)***	(.001)***
Share: Some High School	-.063	-.002	.003	-.092	-.005	.002
	(.019)***	(.002)	(.001)**	(.021)***	(.002)**	(.001)
Share: White-collar occ.	.060	.006	-.002	.057	.004	-.002
	(.010)***	(.001)***	(.0007)**	(.012)***	(.001)***	(.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 6 for separations, Table 7 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.

References

Arbache, Jorge Saba, Andy Dickerson, and Francis Green, "Trade Liberalisation and Wages in Developing Countries," *Economic Journal*, February 2004, *114* (493), F73–96.

Balassa, Bela, "Trade Liberalization and Revealed Comparative Advantage," *Manchester School of Economic and Social Studies*, May 1965, *33*, 99–123.

Bernard, Andrew B., Jonathan Eaton, J. Bradford Jensen, and Samuel Kortum, "Plants and Productivity in International Trade," *American Economic Review*, September 2003, *93* (4), 1268–90.

_ , Stephen J. Redding, and Peter K. Schott, "Comparative Advantage and Heterogeneous Firms," *Review of Economic Studies*, January 2007, *74* (1), 31–66.

Biscourp, Pierre and Francis Kramarz, "Employment, skill structure and international trade: Firm-level evidence for France," *Journal of International Economics*, 2007. Forthcoming.

Brecher, Richard A., "Minimum Wage Rates and the Pure Theory of International Trade," *Quarterly Journal of Economics*, February 1974, *88* (1), 98–116.

Bustos, Paula, "The Impact of Trade on Technology and Skill Upgrading: Evidence from Argentina," November 2005. Universitat Pompeu Fabra, unpublished manuscript.

Cunat, Alejandro and Marc J. Melitz, "Volatility, Labor Market Flexibility, and the Pattern of Comparative Advantage," February 2006. University of Essex, unpublished manuscript.

Davidson, Carl, Lawrence Martin, and Steven Matusz, "Trade and Search Generated Unemployment," *Journal of International Economics*, August 1999, *48* (2), 271–99.

Davis, Steven J., John C. Haltiwanger, and Scott Schuh, *Job creation and destruction*, Cambridge and London: MIT Press, 1996.

Goldberg, Pinelopi Koujianou and Nina Pavcnik, "The Response of the Informal Sector to Trade Liberalization," *Journal of Development Economics*, Special Issue December 2003, *72* (2), 463–96.

Greenaway, David and Richard Kneller, "Firm Heterogeneity, Exporting and Foreign Direct Investment," *Economic Journal*, February 2007, *117* (517), F134–61.

Haltiwanger, John C., "Measuring and Analyzing Aggregate Fluctuations: The Importance of Building from Microeconomic Evidence," *Federal Reserve Bank of St. Louis Review*, May-June 1997, *79* (3), 55–77.

_ , **Adriana Kugler, Maurice Kugler, Alejandro Micco, and Carmen Pagés**, "Effects of Tariffs and Real Exchange Rates on Job Reallocation: Evidence from Latin America," *Journal of Policy Reform*, Special Issue December 2004, *7* (4), 191–208.

Hay, Donald A., "The Post-1990 Brazilian Trade Liberalisation and the Performance of Large Manufacturing Firms: Productivity, Market Share and Profits," *Economic Journal*, July 2001, *111* (473), 620–41.

Hungerford, Thomas L., "International Trade, Comparative Advantage and the Incidence of Layoff Unemployment Spells," *Review of Economics and Statistics*, August 1995, *77* (3), 511–21.

Kletzer, Lori G., *Job Loss from Imports: Measuring the Costs* Globalization balance sheet series, Washington, DC: Institute for International Economics, September 2001.

Kruse, Douglas L., "International Trade and the Labor Market Experience of Displaced Workers," *Industrial and Labor Relations Review*, April 1988, *41* (3), 402–17.

Kume, Honório, Guida Piani, and Carlos Frederico Bráz de Souza, "A Política Brasileira de Importação no Período 1987-98: Descrição e Avaliação," in Carlos Henrique Corseuil and Honório Kume, eds., *A abertura comercial brasileira nos anos 1990: Impactos sobre emprego e salários*, Rio de Janeiro: MTE and IPEA, 2003, chapter 1, pp. 9–37.

Levinsohn, James, "Employment Responses to International Liberalization in Chile," *Journal of International Economics*, April 1999, *47* (2), 321–44.

Melitz, Marc J., "The Impact of Trade on Intra-Industry Reallocations and Aggregate Industry Productivity," *Econometrica*, November 2003, *71* (6), 1695–1725.

Menezes-Filho, Naércio Aquino and Marc-Andreas Muendler, "Labor Reallocation in Response to Trade Reform," *CESifo Working Paper*, March 2007, *1936.*

Muendler, Marc-Andreas, "Trade, Technology, and Productivity: A Study of Brazilian Manufacturers, 1986-1998," *CESifo Working Paper*, March 2004, *1148.*

_ , "Trade and Workforce Changeover in Brazil," in Stefan Bender and Julia Lane, eds., *Conference on the Analysis of Firms and Employees: Quantitative and Qualitative Approaches*, NBER Conference Report series, Chicago and London: University of Chicago Press, forthcoming, chapter 9. (NBER Working Paper *12980*).

Mussa, Michael, "Dynamic Adjustment in the Heckscher-Ohlin-Samuelson Model," *Journal of Political Economy*, October 1978, *86* (5), 775–91.

Obstfeld, Maurice and Kenneth Rogoff, *Foundations of international macroeconomics*, Cambridge, MA and London: MIT Press, 1996.

Olley, G. Steven and Ariel Pakes, "The Dynamics of Productivity in the Telecommunications Equipment Industry," *Econometrica*, November 1996, *64* (6), 1263–97.

Paes de Barros, Ricardo and Carlos Henrique Corseuil, "The Impact of Regulations on Brazilian Labor Market Performance," in James J. Heckman and Carmen Pagés, eds., *Law and employment: Lessons from Latin America and the Caribbean*, NBER Conference Report series, Chicago and London: University of Chicago Press, 2004, pp. 273–350.

Pavcnik, Nina, "Trade Liberalization, Exit, and Productivity Improvement: Evidence from Chilean Plants," *Review of Economic Studies*, January 2002, *69* (1), 245–76.

Raith, Michael, "Competition, Risk, and Managerial Incentives," *American Economic Review*, September 2003, *93* (4), 1425–36.

Ramos, Roberto Luís Olinto and Claudia Nessi Zonenschain, "The Performance of the Brazilian Imports and Exports Based on the System of National Accounts: 1980-1998," August 2000. IBGE Rio de Janeiro.

Ribeiro, Eduardo Pontual, Carlos Henrique Corseuil, Daniel Santos, Paulo Furtado, Brunu Amorim, Luciana Servo, and André Souza, "Trade Liberalization, the Exchange Rate and Job Flows in Brazil," *Journal of Policy Reform*, Special Issue December 2004, *7* (4), 209–23.

Rivers, Douglas and Quang H. Vuong, "Limited Information Estimators and Exogeneity Tests for Simultaneous Probit Models," *Journal of Econometrics*, November 1988, *39* (3), 347–66.

Roberts, Mark J., "Employment Flows and Producer Turnover," in Mark J. Roberts and James R. Tybout, eds., *Industrial evolution in developing countries: Micro patterns of turnover, productivity, and market structure*, Oxford and New York: Oxford University Press for the World Bank, 1996, pp. 18–42.

Saint Paul, Gilles, "Is Labour Rigidity Harming Europe's Competitiveness? The Effect of Job Protection on the Pattern of Trade and Welfare," *European Economic Review*, April 1997, *41* (3-5), 499–506.

Santos-Paulino, Amelia and A. P. Thirlwall, "Trade Liberalisation and Economic Performance in Developing Countries: Introduction," *Economic Journal*, February 2004, *114* (493), F1–3.

Tybout, James R., "Plant- and Firm-level Evidence on "New" Trade Theories," in James Harrigan and E. Kwan Choi, eds., *Handbook of International Trade*, Vol. 1 of *Blackwell Handbooks in Economics*, New York: Blackwell Publishers, July 2003, chapter 13, pp. 388–415.

Verhoogen, Eric A., "Trade, Quality Upgrading and Wage Inequality in the Mexican Manufacturing Sector," February 2007. Columbia University, unpublished manuscript.

Wacziarg, Romain and Jessica Seddon Wallack, "Trade Liberalization and Intersectoral Labor Movements," *Journal of International Economics*, December 2004, *64* (2), 411–39.

Winters, L. Alan, "Trade Liberalisation and Economic Performance: An Overview," *Economic Journal*, February 2004, *114* (493), F4–21.