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Labor Reallocation in Response to Trade Reform

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Labor Reallocation in Response to Trade Reform*

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Abstract

Tracking individual workers across employers and industries after Brazil's trade liberalization in the 1990s shows that foreign import penetration and tariff reductions trigger worker displacements but that neither comparative-advantage industries nor exporters absorb displaced workers for years. There are significantly more displacements and fewer accessions in comparative-advantage industries and at exporters. These findings are robust to instrumenting trade barriers and export status with product demand at Brazil's export destinations and real exchange rate components. Worker effects are important predictors of labor turnover. Trade liberalization is associated with significantly more transitions to informal work status and self-employment. Output is reallocated to more productive firms after trade reform but, given fast labor-productivity growth, this product reallocation is not accompanied by similar labor reallocation.

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

JEL Classification: F14, F16, J23, J63

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1 Introduction

Economists have long studied the consequences of international trade. Numerous empirical studies investigate the impact of trade on economic outcomes at the level of the plant, firm or the country as a whole. Yet research to examine the impact of trade on workers' individual employment trajectories is scant.[1] Both classic and firm-level trade theories predict that an essential aspect of trade integration is worker reallocation, either across sectors according to comparative advantage (Ricardo, Heckscher-Ohlin-Samuelson) or across firms depending on export status (Bernard et al. 2003, Melitz 2003).

Trade liberalization episodes in developing countries are natural environments to investigate consequences of trade integration. For numerous Latin American economies, empirical studies present evidence of considerable productivity advances at the level of plants or firms in response to trade liberalization and other pro-competitive reforms. But these firm-level gains hardly translate into GDP growth in Latin America, where aggregate economic performance in the 1990s is widely perceived as disappointing in the wake of reforms designed to boost growth.[2] Some researchers attribute low growth measures to a mismeasured price deflator following trade reform (Chamon and Carvalho Filho 2006). Other commentators attribute the lacking economic performance to limited coordination among initiatives, and to unbalanced reforms neglecting labor-market flexibility (e.g. Singh et al. 2005). No empirical study thus far has used individual-level data to support the latter claim.

We use, to our knowledge for the first time in the context of trade reform, economy-wide linked employer-employee data to follow Brazilian workers across employers and industries before and after a period of rapid trade liberalization. Published research to date employs more aggregate data, at the level of plants, firms, sectors, or countries. Much of the preceding literature uses Davis and

[1]Kletzer (2001) is a notable exception that we discuss below. Tybout (2003) surveys firm- and plant-level studies, Winters (2003) summarizes evidence on trade and development.

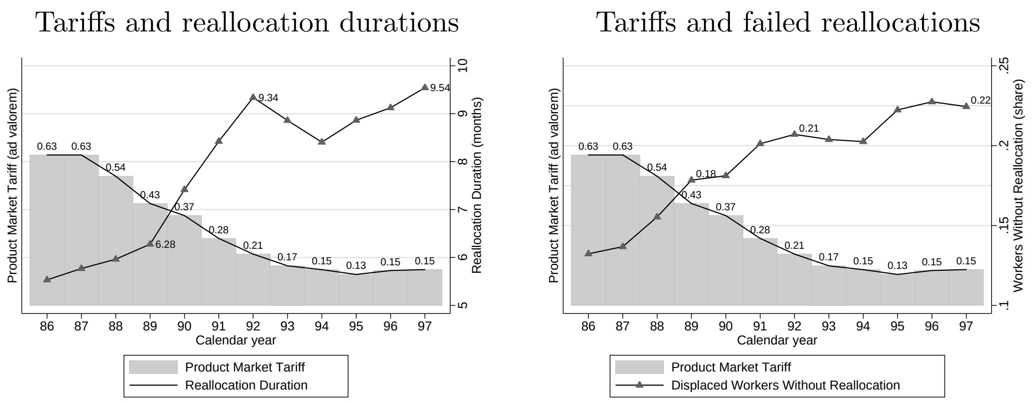
[2]The World Bank (2005) discusses economic performance of developing countries in its document on reform during the 1990s. Bosworth and Collins (2003) report for Latin America an average annual output-per-worker growth rate of .9 percent in the 1990s. Among the micro-level studies of trade reforms, Eslava et al. (2004) report estimates that imply an annual TFP growth rate of 3.5 percent in Colombian manufacturing plants in the 1990s, and Schor (2004) reports estimates that imply an average annual TFP growth rate of 1.9 percent at Brazilian manufacturing firms from 1990 to 1998. La Porta and Lopez de Silanes (1999) document productivity advances after privatization in Mexico between 1983 and 1991, and Pavcnik (2002) reports firm-level productivity gains following Chile's early trade reform in the region.

Haltiwanger (1992) turnover measures of job creation, destruction, and sectoral churning to examine labor-market effects of trade liberalization.[3] Churning is the excess job turnover beyond net employment change. For a panel of six Latin American countries, Haltiwanger et al. (2004) find that tariff reductions are associated with heightened within-sector churning and net employment reductions at the sector level.[4] Biscourp and Kramarz (2007) use detailed information on trade at the firm-level for France and detect a statistically significant association between imports and job destruction. Protection typically benefits sectors with a comparative disadvantage. So, the employment contraction in formerly protected industries or at importing firms, as documented in those studies, is both expected and desirable: trade theory welcomes factor displacements from activities with a comparative disadvantage. But does a successful reallocation to comparative-advantage industries or exporting firms ensue? It takes linked employer-employee data to scrutinize the success of worker reallocations across firms and industries after displacement.

We combine information from numerous sources and construct a comprehensive data set to test labor-market predictions of trade theories. From a worker register at Brazil's labor ministry we gather detailed information on education, occupation, age, tenure, labor-market experience and employment status for every formal-sector worker and the worker's identified employer. At the sector level, we construct measures of comparative advantage and link information on tariffs, import penetration, sectoral real exchange rates, industrial concentration, and foreign direct investment. At the plant- and firm-level, we obtain time-varying information on workforce size and composition, private or public ownership, labor

[3] Examples include Davis et al. (1996) for the U.S., Roberts (1996) for developing countries, and Ribeiro et al. (2004) for Brazil.

[4] Earlier studies on trade reform do not identify significant effects. Roberts (1996) reports no clear effect of trade exposure on employment changes at plants in Chile and Colombia, once 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 in the presence of concomitant macroeconomic shocks. Neither do Davis et al. (1996) identify a clear effect of trade on factor reallocation using U.S. data, or do Wacziarg and Wallack (2004) detect statistically significant labor reallocations in a cross-country study of trade-liberalization periods. In our Brazilian data, sector variables are not statistically significant predictors of net employment changes at the plant level (Muendler 2007), whereas worker-level regressions on the same data in this paper reveal salient correlations between sector variables and labor turnover at the one-percent significance level. This suggests that unobserved workforce heterogeneity may hamper regressions at more aggregate levels. Studies considering exchange rate effects, however, do find systematic effects on employment flows in France and the U.S. (Gourinchas 1999, Klein et al. 2003).



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

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

productivity, and export status. The time dimension of up to 16 years between 1986 and 2001 allows us to estimate worker-fixed effects that control for non-random matches of unobserved worker types to employers.

We document for the case of Brazil's large-scale trade reform that incomplete labor reallocation is a reason for the country's weak aggregate economic performance. Our linked employer-employee data show that increasing import penetration and sectoral differences in tariff reductions predict significant increases in worker separations. But displaced formal-sector workers are not being reallocated for extended periods of time. Figure 1 illustrates this fact by plotting *ad valorem* product-market tariffs against the evolution of reallocation durations in months (on the left) and the share of displaced workers with no formal-sector reallocation for four years (on the right). The curves follow opposite trends. Brazil's tentative tariff reductions prior to 1990 prove little effective because of binding non-tariff barriers, whereas the complete removal of non-tariff barriers on January 1, 1990 unleashes the full pro-competitive effect of the prior tariff cuts and those following between 1990 and 1993. Worker reallocations take six months or less prior to 1990, if successful. By 1993, durations of successful reallocations are more than nine months and never fall back below eight months again. The share of displaced-worker reallocations that are to fail for four years or more, rises from 18 percent and less before 1990 to 21 percent by 1992 and to 22 percent by 1997.

Note that, for the pool of displaced workers with no reallocation for four years, trade reform on January 1, 1990 matters as early as 1986.

Our results differ markedly from previous research in the area. Beyond sectoral labor-turnover measures, we assess the direction of labor shifts across industries and firms to test predictions of classic and heterogeneous-firm trade theories. Contrary to those predictions, worker-level regressions using our linked employer-employee data demonstrate that employers in comparative-advantage industries and exporters displace workers significantly more frequently. Exporters hire workers significantly less frequently than the average firm, and employers in comparative-advantage industries hire workers significantly less frequently conditional on the interactions between comparative advantage, export status and tariffs. So there are marked net employment shifts out of comparative-advantage industries and away from exporters. Worker-specific effects are statistically important components of the reallocation prediction.

In a pioneering study of trade effects on employment trajectories, Kletzer (2001) compares the labor-market experience of workers displaced from import-competing U.S. manufacturing industries to displaced U.S. workers from other sectors and detects no marked differences in the workers' subsequent employment experience. Estimating a hazard model for the reallocation duration of displaced workers in our formal-sector data shows that successful reallocations to an exporter are significantly less frequent than to an average firm. But reallocation hazards also show that displacement from an exporter predicts a significantly shorter reallocation duration, iff successful. Beyond import competition, our analysis accounts for the differential effects of comparative advantage and tariffs across sectors and the firm-level heterogeneity across employers within sectors. These distinctions are important. Though our results refute the prediction that comparative-advantage industries and exporters expand net employment, interactions show that, within comparative-advantage industries, exporters hire significantly more workers than nonexporters and tariff cuts heighten the exporter-nonexporter difference. Those findings are consistent with predictions of heterogeneous-firm trade theory: within comparative-advantage industries, exporters hire workers significantly more frequently than non-exporters.

Regressions for individual household members from a separate household-survey data source show that reduced tariffs and heightened import penetration predict significantly more transitions into informality and unemployment, and fewer transitions from informality back to formal employment at annual horizons. Goldberg and Pavcnik (2003), in contrast, report no effect of trade reform on the incidence of informal work using sectoral data for Brazil. A possible reason for the difference in results is that the household survey allows us to follow the same individual for up to sixteen months so that we can control for prior labor-market experience and observe simultaneous industry and work status transitions. In summary, trade liberalization is associated with longer durations of formal-job reallocations, more frequent transitions to informal work status and unemployment, and more frequent failures of formal-job reallocations for several years.

This labor-market evidence stands in a seeming contrast with the common finding that product-market shares are reallocated to more productive firms and exporters (see Tybout 2003). If labor productivity rises faster than output, however, then product-market reallocations to more productive firms and simultaneous workforce shifts away from more productive firms are a theoretical possibility. We document for firms in Brazil's manufacturing survey that this theoretical possibility is Brazil's reality during the 1990s: additional firm-level output is associated with less employment while labor productivity advances fast. Worker-level regressions for the period following trade reform show that firms with higher labor productivity hire significantly fewer workers than the average firm, and exporters hire significantly fewer workers given labor productivity. These results suggest that extensions of trade models to include endogenous firm-level productivity change in response to trade can account for empirical facts that classic and current heterogeneous-firm models do not capture.

We perform several robustness checks and rule out alternative hypotheses. We address the potential simultaneity of trade policies, exporting status and labor turnover by predicting tariffs, import penetration and export status with instrumental variables for export demand by sector and year (using imports from other source-countries than Brazil in seven foreign destination regions) and with exogenous components of the sectoral real exchange rate. Tests show the instruments to be strong predictors, and linear fixed-effects regressions corroborate our main hypotheses that firms in comparative-advantage sectors and exporters separate from their workers significantly more frequently and that exporters hire significantly less frequently than the average firm. To our surprise, we find the trade effects on separations and accessions to hardly vary by worker skill group and conclude that skill-biased technical change does not significantly interact with our findings. We investigate the relevance of concomitant economic changes during the sample period, including a trend shift of employment from manufacturing to services, policy changes such as macroeconomic stabilization, surging foreign direct investment inflows, privatization of state-owned companies, the intensified outsourcing of service jobs, and the constitutional reform of labor-market regulations preceding trade reform. We condition on year effects in all regressions and find separation rates in manufacturing to monotonically rise over the sample period and accession rates to monotonically drop—consistent with Brazil's increasingly overvalued exchange rate during the sample period and with a trend shift of employment out of manufacturing. Conditioning on foreign direct investment inflows at the sector level, on privatization at the plant level, on the susceptibility of a job to outsourcing at the occupation level, and on industry effects to capture sectoral labor-turnover interactions with preceding labor-market reform, allows us to corroborate our main hypotheses: employers in comparative-advantage industries and exporters separate from workers significantly more frequently, and exporters hire workers significantly less frequently than the average firm.

The paper proceeds 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 presents estimates of work status transitions from a household perspective. Section 5 analyzes worker separations, accessions and reallocation durations and identifies sector and firm predictors that explain the observed delays and failures in the reallocation process. Section 6 subjects these predictions to numerous robustness checks. Section 7 discusses the findings and their implications for trade theory. Section 8 concludes.

2 Data

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

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

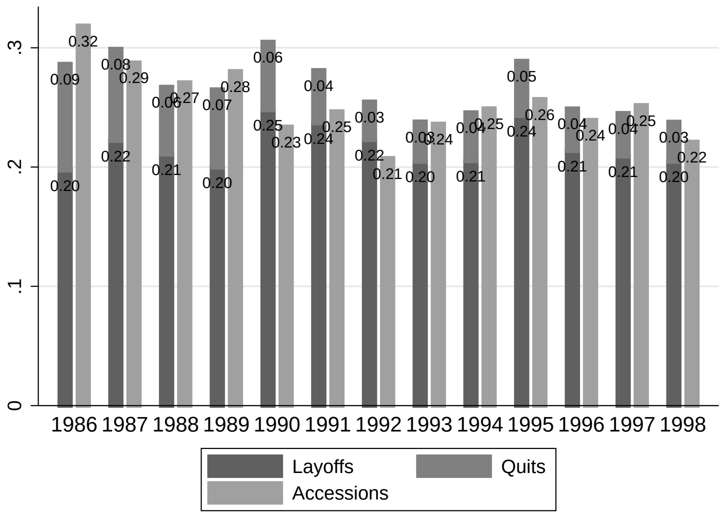
Linked employer-employee data. *RAIS* (*Relação Anual de Informações Sociais*) provides comprehensive annual information on workers formally employed in any sector (including the public sector). Our full data include 71.1 million workers with 556.3 million job spells at 5.52 million plants in 3.75 million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

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

We only keep prime-age male workers, 25 to 64 years old, in the worker sample in order to focus on workers after their first labor-force entry and to be little affected by labor supply changes. For most separation statistics, we remove mul-

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

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



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

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

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

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

Table 1 shows in the upper panel that the changes to gross separation and accession rates in manufacturing result in a net labor-force shift out of manufacturing over the course of the 1990s. While the manufacturing sector employs 26 percent of Brazil's national formal workforce in 1990, the share is only 21 percent by 1998. The primary sector (agriculture and mining), commerce and services employ larger shares of formal workers in 1998 than in 1990. Construction, utilities, and public administration are shown as other sectors in Table 1 and exhibit a reduction in relative importance between 1990 and 1998. Our sample represents a population of around 10.8 million prime-age male workers in Brazil in 1990 and 11.6 million in 1998. The middle panel of Table 1 presents employment shares for the metropolitan areas of Brazil and shows that metropolitan labor markets exhibit employment shares and undergo changes in relative employment similar to the economy as a whole. More than half of Brazil's total formal-sector employment is in metropolitan areas.

Metropolitan household data. The Brazilian monthly employment survey *PME* (*Pesquisa Mensal de Emprego*) provides details on work status. *PME* data derive from a random sample of households in six metropolitan areas (São Paulo, Rio de Janeiro, Belo Horizonte, Porto Alegre, Salvador, Recife). The data are collected by Brazil's statistical bureau IBGE from a rotating panel similar to the U.S. *PSID*. *PME* follows households for 16 months, with an eight-month interval after the fourth interview.[8] Changes to the sample design adversely affect worker panels starting in odd years. So, we use only individuals whose first survey occurs in 1986, 1988, 1990, 1992, 1994, 1996 or 1998. As with *RAIS*, we restrict our sample to prime-age male workers. We trace work-status changes 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

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

Table 1: EMPLOYMENT BY SECTOR AND FORMALITY STATUS

	Primary	Manuf.	Comm.	Services	Other	*Total*[a]
	(1)	(2)	(3)	(4)	(5)	(6)
			Allocation nationwide (*RAIS*)			
1990	.029	.263	.111	.284	.314	10,763
1998	.064	.207	.134	.308	.286	11,640
			Allocation in metropolitan areas (*RAIS*)			
1990	.015	.270	.104	.309	.302	5,965
1998	.023	.198	.125	.369	.285	6,057
			Informality share in metropolitan areas (*PME*)			
1990	.159	.063	.109	.117	.298	
1998	.232	.120	.154	.169	.341	

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

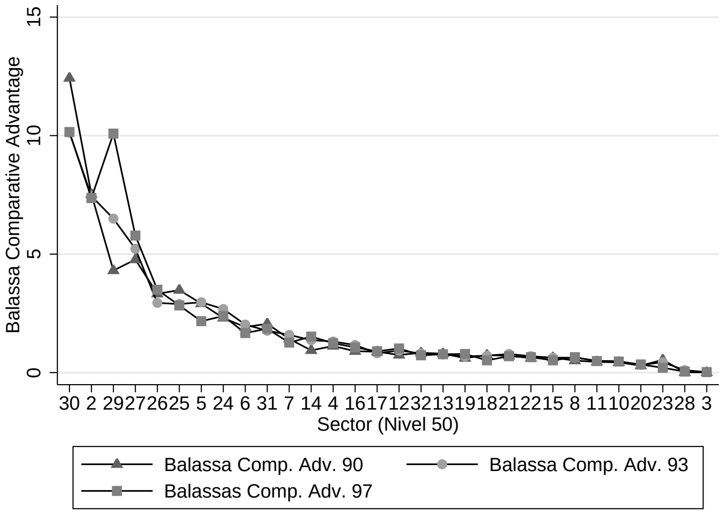
Sources: *RAIS* 1990 and 1998, male workers nationwide (1% random sample) and in metropolitan areas (5% random sample), 25 to 64 years old, and employed on December 31; and *PME* 1990 and 1998, male workers 25 to 64 years old, and employed at September interview. Primary sector includes agriculture and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

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

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

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

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



Source: UN Comtrade 1986-98. Sectors at *Nível 50* ranked by sector-fixed Balassa comparative advantage (for sector codes see Table 16 in the Appendix). Estimates of sector-fixed Balassa comparative advantage from regression of Balassa advantage on sector indicators, year indicators, output tariffs and input tariffs between 1990 and 1998.

Figure 3: **Balassa Comparative Advantage**

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

Brazil's comparative advantage pattern is remarkably stable over the sample period. Figure 3 ranks manufacturing industries by their sector-fixed component 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 the regression; and year indicators are neither individually nor jointly different from zero at common significance levels.[9] Figure 3 illustrates the regressions results. With the exception of processed sugar (sector 29), Brazil's comparative advantage changes hardly at all. Removal of the sugar-processing sector from our regressions shows that results are not sensitive to its inclusion. Overall, the sector ranking by comparative advantage remains

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

largely unaltered over time.

Our main instrumental variables for firm-level 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: 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).[10] Additional instruments are components of the sectoral real exchange rate: the U.S. dollar exchange rate to Brazilian currency, rebased to Brazilian Real in August 1994, and sector price levels in the U.S. and the EU. We relegate details on these instruments and a description of additional sector variables, including tariff and market-penetration measures, to the Appendix.

Firm data. We combine the linked employer-employee data from *RAIS* with additional firm-level data.[11] We use annual customs office records from *SECEX* (*Secretaria de Comércio Exterior*) on exports 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 link firm-level labor productivity measures from the manufacturing survey *PIA* to *RAIS*. Details are discussed in the Appendix.

[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. 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.

[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 random subsample of (*PIA*) firms. Our regressions control for plant employment, so exports per worker would not add information. The binary export-status indicator is found to be a highly significant predictor of worker turnover in regressions. In fact, the export indicator predicts turnover patterns similar to sector-level comparative advantage, conditional on sector, plant, worker and job characteristics.

Table 2: Productivity Variation Across Firms and Over Time

	TFP and Output shares				Labor Prod. and Employment shares			
	Cross section			Ann. chg.	Cross section			Ann. chg.
	wgtd.	unwgtd.	cov.	raw cov.[a]	wgtd.	unwgtd.	cov.	raw cov.[a]
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
1986	1.018	.924	.095		1.011	1.019	-.008	
1990	1.000	.899	.101	.065	1.000	.997	.003	-.029
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 average of raw covariance 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): $\underline{y_t} = \bar{y}_t + \sum_i \overline{\Delta}\theta_{it}\overline{\Delta}y_{it}$, where $\underline{y_t}$ 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).

3 Labor Reallocation and Trade

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

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

Table 2 decomposes total factor productivity (columns 1-4) and labor productivity (columns 5-8) into the contributions of firm-level productivity and firm-level weights, where the weights are output in the case of total factor productivity and employment in the case of labor productivity. The statistics are based on output and employment at formal-sector manufacturing firms. Following Olley and

Pakes (1996), aggregate productivity in the cross section of firms (columns 1 and 5) is split into the unweighted mean productivity level (columns 2 and 6) and the covariance between deviations of the weights and productivities from annual means (columns 3 and 7). The overall TFP gain of 3.5 percent between 1990 and 1998 is modest (column 1).[13] Substantial capital accumulation contributes to the faster increase in labor productivity by 7.3 percent between 1990 and 1998 (column 5). Alongside, Table 2 reports the raw covariance of year-over-year productivity changes at surviving firms (columns 4 and 8)—a term in the Haltiwanger (1997) decomposition over time.[14]

The decompositions in Table 2 show for the cross section of Brazilian manufacturers that firms with higher total factor productivity (TFP) do command larger output shares (column 3), and that TFP improvements among survivors are associated with gains in output shares (column 4). The cross-sectional covariance between labor productivity and employment shares, however, 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).[15] So, firms with increasing productivity expand output shares while reducing employment.

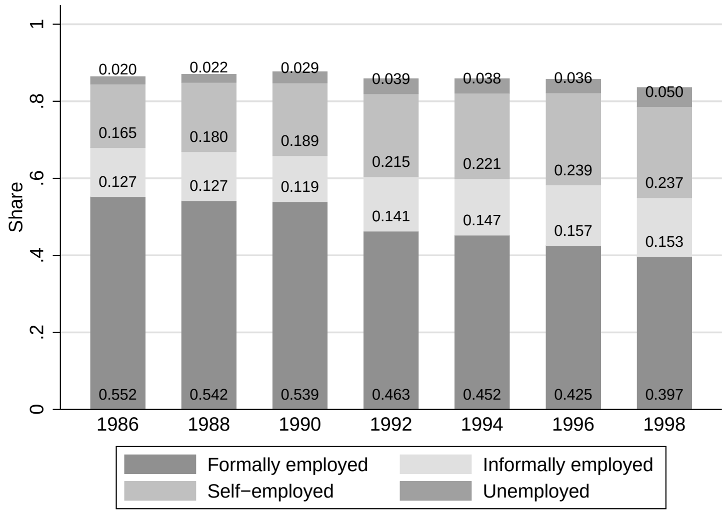
Household survey data offer a perspective on labor reallocation beyond the formal sector. Figure 4 documents for prime-age male workers that formal work status drops from a share of 54 percent to 40 percent between 1990 and 1998 across all sectors in metropolitan areas, while informal employment increases from 12 to 15 percent. Over the same period, self employment increases from 20 to 24 percent, unemployment from 3 to 5 percent, and withdrawals from the labor force rise from 12 to 16 percent.

Economic reforms. In 1990, the Brazilian government breaks with the country's decade-old import substitution policy and embarks on drastic trade liberalization. Contrary to tentative *ad valorem* tariff reductions during the late 1980s—rendered largely ineffective because of binding non-tariff barriers (Kume,

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

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

[15]Mostly firm exits raise the covariance between labor productivity and employment in the cross section over time (column 7).



Source: *PME* 1986-98, male workers, 25 years or older and employed in metropolitan area. Remaining share: withdrawn from labor force.

Figure 4: **Work status shares in metropolitan areas**

Piani and Souza 2000)—, 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 in the Introduction documents the drop in product tariffs from an average level of 63 percent in 1987 to 15 percent by 1997. The new tariff structure also reduces the cross-sectoral dispersion. While product tariffs range between 21 (metallic products) and 63 percent (apparel and textiles) in 1990, they drop to a range spanning between 9 percent (chemicals) and 34 percent (transport equipment) in 1997. Manufacturing industries receive effective protection in both years. In 1990, product tariffs are around 45 percent above intermediate-input tariffs in value-added terms. By 1997, however, the reduced cross-sector dispersion of tariffs results in a smaller rate of effective protection of about 20 percent on average.

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

Workforce characteristics and trade exposure. Table 3 provides a summary comparison of workforce characteristics across 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. Unconditional means conceal substantial employer and workforce heterogeneity, however. The average exporter (that employs a prime-age male worker) is active in a sector with a slightly lower than average comparative advantage level. Similarly, fewer prime-age male workers are employed at exporters in a top comparative-advantage sector than at exporters overall. The reason is that export volumes per employee at exporters in a top comparative-advantage sector exceed export volumes from firms in a sector without comparative advantage.

Expectedly for a country with a history of import-substitution industrialization, Brazil's top comparative-advantage industries have lower-than-average tariffs and import-penetration rates. The top comparative-advantage industries attract larger foreign-direct investment inflows, are slightly less concentrated than the average sector, and employ slightly fewer workers at state-owned companies. Plants in top comparative-advantage industries and exporters have larger workforces than average (85 and 326 workers more, respectively, than the average formal-sector manufacturing plant with 257 workers). Recall that our sample is a random draw of workers from the formal-sector worker population so that larger plants are more likely to be included. Labor productivity at exporters is slightly higher than average, and at firms in comparative-advantage industries somewhat

Table 3: *RAIS* Summary Statistics for Manufacturing

	All sectors and firms		5th comp. adv. quintile	Exporter
	Mean	Std.Dev.	Mean	Mean
	(1)	(2)	(3)	(4)
Outcomes				
Indic.: Separation	.282	.450	.314	.260
of these: 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
Sector-level covariates				
FDI Flow (USD billion)	.110	.334	.263	.103
Herfindahl Index (sales)	.089	.056	.083	.098
Share: Jobs at private firms	.955	.019	.966	.955
Plant-level covariates				
Log Employment	5.148	1.952	5.551	6.210
Log Labor Productivity	36.315	6.679	36.449	36.331
Worker-level covariates				
Tenure at plant (in years)	.952	1.208	.778	1.248
Pot. labor force experience	25.276	9.971	26.115	25.155
Middle School or less	.785	.411	.854	.745
Some High School	.151	.358	.108	.171
Some College	.020	.141	.012	.028
College Degree	.038	.191	.021	.052
Prof. or Manag'l. Occ.	.085	.278	.069	.102
Tech'l. or Superv. Occ.	.082	.275	.061	.098
Unskilled Wh. Collar Occ.	.070	.255	.079	.075
Skilled Bl. Collar Occ.	.636	.481	.646	.623
Unskilled Bl. Collar Occ.	.102	.303	.120	.088
Indic.: Outsourceable job	.252	.434	.234	.294

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 (146,800 observations in separation, 112,971 in accession sample). Sector information at subsector IBGE level.

Table 4: ANNUAL SECTOR TRANSITIONS AND FAILED RE-ACCESSIONS

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

[a]Balassa (1965) comparative advantage, transition year quintile (5th: strongest advantage).
Source: *RAIS* 1986-2001 (1% random sample), male workers nationwide, 25 to 64 years old. UN Comtrade 1986 for Balassa comparative advantage; defined at two-digit sector level (Subsector IBGE). Transition frequencies are job accessions in Brazil within one year after separation, based on last employment of year (highest paying job if many). Failed accessions are separations followed by no formal-sector accessions anywhere in Brazil within a year, excluding workers with prior retirement or death, or age 65 or above on earlier job.

higher still.

In line with the higher-than-average turnover in comparative-advantage industries, prime-age male workers' tenure at plants in comparative-advantage industries is lower than average. Similarly, lower labor turnover at exporters is associated with longer-than-average worker tenure. In comparative-advantage industries, workers are slightly older than average while somewhat younger than average at exporters. Workers employed in comparative-advantage industries are less educated than average but exporter employees have higher-than-average schooling. Similarly, jobs in comparative-advantage industries are less skill intensive than average but jobs at exporters exhibit higher-than-average skill intensity. Jobs in comparative-advantage industries are less susceptible to outsourcing, but more susceptible at exporters. Overall, these remarkable mean differences between the average exporter and the average firm in a top comparative-advantage sector suggest that substantial employer heterogeneity prevails in top comparative-advantage industries with diverse nonexporters shifting the mean characteristics.

Worker reallocation. Table 4 reports transitions of displaced prime-age male workers from formal-sector jobs to other formal-sector jobs at the annual horizon (columns 1-6) for the period 1986-2001, and the share of displaced workers with no observed formal-sector rehiring within a year (column 7). Retained workers do not enter the statistics. Agricultural, mining and manufacturing plants are grouped into their sector's comparative advantage quintiles at the subsector IBGE level.[16] All other sectors—commerce, services, construction, utilities, and public administration—are considered nontraded for the purposes of Table 4.

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

The dominant fraction of workers with displacement from a traded-goods plant, about a third, finds employment in nontraded-output sectors (column 6). And almost as many workers with displacements from a traded-goods sector, roughly another third, are not rehired into any formal job within a year (column 7). Three out of four workers who are not reallocated at the annual horizon, but who find reemployment in subsequent years, move to the nontraded sector (second-to-last row) and, among the traded industries, mostly into high-quintile industries. A repetition of the statistical exercise for different subperiods during the 1986-98 period shows that transition patterns within traded-goods industries remain remarkably stable. The overall shares of re-accessions at nontraded-output sectors and of failed annual re-absorptions drop by 3 and 1 percent between 1990-

[16] 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 striking difference.

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.

91 and 1996-97. These patterns are consistent with the idea that work status changes out of formality (recorded as failures here), and jobs in nontraded-output sectors, provide a buffer for labor reallocation after trade reform.

Linked employer-employee data track worker reallocations at additional margins. Table 5 shows flows of prime-age male workers between nonexporters and exporters for the two periods 1990-91 and 1996-97. Around a million prime-age male workers are successfully reallocated each period, and 91 percent of them shift to nonexporters while 9 percent transition to exporters. The share of former exporter workers who are rehired at exporters (23 percent in both periods) exceeds the share of former nonexporter workers with a reallocation to exporters (7 percent in both periods). The small magnitude of transitions to exporters, with less than one in ten displaced workers shifting to an exporter overall, and the rehiring bias at exporters towards former exporter workers suggest that labor shifts from nonexporters to exporters are not a major channel of worker reallocation.[17]

Descriptive evidence in this section is based on unconditional means. The remainder of the paper subjects reallocation statistics to multivariate controls. Reallocations to the nontraded-output sector, for instance, can be partly due to an overvalued real exchange rate and time trends that favor services expansion. Concomitant reforms and firm and worker heterogeneity require attention. The following two sections analyze two main aspects of the observed employment

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

shifts. The next section investigates reallocations to work status outside formal employment. The section after the next analyzes predictors of increased separations, which fill the pool of workers to be reallocated and delay the average worker in the pool, and predictors of reduced accessions in the formal sector, which do not contribute to emptying the pool and thus also delay the average worker awaiting formal-sector reallocation.

4 Work Status Transitions

To investigate how Brazil's varying trade exposure predicts transitions between types of work status for individual workers, we estimate a multinomial logit model using *PME* household survey data.[18] *PME* reports a single work status transition per identified household member at the annual horizon.

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

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

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

[18]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 6: Work Status Transitions from Formal Employment

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

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

covariates $\mathbf{z}_{S(i),t}$ are well defined, but do not impose a sector restriction on job observations at $t+1$.

Transitions out of formality. The set of work status outcomes for a worker with a formal manufacturing job contains five alternatives: (1) the worker retains the formal manufacturing job or switches to a new formal job (not necessarily in manufacturing); (2) the worker moves to an informal job (not necessarily in manufacturing); (3) the worker moves to self-employment (not necessarily in manufacturing); (4) the worker moves to unemployment; and (5) the worker withdraws from the labor force. To capture the effect of changing tariff dispersions for a sector's effective rate of protection, we include product tariffs and intermediate-input tariffs in the multinomial logit regressions. Table 6 presents the predictions.

Elevated product-market tariffs are associated with significantly lower odds of transitions out of formal manufacturing employment and into informality or self employment. The level of product-market tariffs is not a statistically significant predictor of unemployment but is associated with significantly higher odds of withdrawals from the labor force. Intermediate-input tariff coefficients show converse signs, and predict significantly more transitions out of formality and into informality. The sign reversals are consistent with the notion that elevated intermediate input tariffs aggravate competitive pressure in product markets, whereas high product-tariff barriers reduce competitive pressure. As described above, Brazil's manufacturing industries face a drop in the effective rate of protection, that is a faster decline in product tariffs than in input tariffs. Together, coefficient magnitudes and the relatively faster drop of product-market tariffs imply that Brazil's trade liberalization predicts more transitions out of formal manufacturing employment and into informality or self-employment. This worker-level evidence on transitions into informality challenges findings in Goldberg and Pavcnik (2003), who report no significant effect of trade liberalization on the incidence of informality in sector data for Brazil.

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

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

Elevated product-market tariffs are associated with significantly higher odds of transitions out of informality in manufacturing and into formality or self employment. Lower product-market tariffs are also associated with significantly lower odds of a transition from informality into unemployment or withdrawals from the labor force. Intermediate-input tariff coefficients, however, are not different from zero at common significance levels. Workers with continuously reported informal-sector employment during the first four months of observation are significantly less likely to leave informality status over the following year. As for household

Table 7: Work Status Transitions from Informal Employment

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

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

members with initial formal work, longer labor force experience predicts more transitions into self-employment and fewer into unemployment. Higher educational attainment predicts significantly less transitions out of informality and into formal work, self employment or unemployment but has no significant effect on labor force withdrawals.

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

5 Separations, Accessions, and Reallocation

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

5.1 Separations and accessions

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

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

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

Table 8 presents conditional logit estimates with worker-fixed effects for separations from formal manufacturing jobs. Separations are significantly more frequent in sectors with a stronger comparative advantage and at exporters. Elevated product tariffs predict lower separation rates from formal jobs (though only

Table 8: Conditional Logit Estimation of Separations

	(1)	(2)	(3)	(4)	(5)
Balassa (1965) Comp. Adv.	.079				.169
	(.021)***				(.024)***
Exporter Status		.289			.283
		(.028)***			(.028)***
Product Market Tariff			-.114		-.715
			(.416)		(.426)*
Intm. Input Tariff			1.621		2.901
			(.633)**		(.678)***
Import Penetration				.773	1.252
				(.353)**	(.388)***
Sector-level covariates					
Sector real exch. rate (EP^*/P)	.725	.835	.345	.693	-.404
	(.624)	(.626)	(.640)	(.631)	(.645)
FDI Flow (USD billion)	-.025	-.012	-.017	-.013	-.048
	(.020)	(.020)	(.020)	(.020)	(.020)**
Herfindahl Index (sales)	-.369	-.514	-.394	-.653	-.347
	(.317)	(.316)	(.329)	(.325)**	(.343)
Plant-level covariates					
Log Employment	-.343	-.370	-.341	-.339	-.377
	(.011)***	(.011)***	(.011)***	(.011)***	(.011)***
Share: Middle School or less	-.745	-.653	-.714	-.711	-.658
	(.131)***	(.131)***	(.131)***	(.131)***	(.132)***
Share: Some High School	-.439	-.387	-.435	-.438	-.388
	(.148)***	(.148)***	(.147)***	(.147)***	(.148)***
Share: White-collar occ.	.725	.704	.742	.742	.694
	(.075)***	(.074)***	(.074)***	(.074)***	(.075)***
Worker-level covariates					
Tenure at plant (in years)	1.367	1.350	1.362	1.363	1.351
	(.036)***	(.036)***	(.036)***	(.036)***	(.036)***
Pot. labor force experience	.006	.006	.006	.006	.006
	(.002)**	(.002)**	(.002)**	(.002)**	(.002)**
Unskilled Wh. Collar Occ.	-.256	-.252	-.260	-.256	-.263
	(.067)***	(.067)***	(.067)***	(.067)***	(.067)***
Obs.	145,417	145,417	145,417	145,417	145,417
Pseudo R^2	.148	.149	.148	.148	.150

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. Controlling for year effects. 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.

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, heightened import penetration predicts significantly higher displacement odds. We include observed market penetration with imports to proxy for the pro-competitive effect of reduced non-tariff barriers. Point estimates and statistical significance of these coefficients are hardly affected as the specification is gradually enriched (moving from column 1 to column 5). FDI inflows into the sector predict a statistically significant reduction in displacement rates. The sectoral real exchange and the Herfindahl concentration index have no significant predictive power after conditioning on year effects.

Before discussing plant and worker-level variables, we turn to the opposite margin. Table 8 presents conditional logit estimates of accessions into formal manufacturing jobs, conditional on worker-fixed effects. Mirroring the signs from separation regressions, accession rates are lower in sectors with stronger comparative advantage, when we condition on other trade-related variables (column 5). The coefficient is not statistically significant at conventional levels in this regression (but will become statistically significant when controlling for higher-order interactions between trade variables in Table 15). 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.

Larger manufacturing plants exhibit less turnover: they displace significantly fewer (Table 8) and they hire significantly fewer workers (Table 9). 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, consistent with a trend shift of

Table 9: Conditional Logit Estimation of Accessions

	(1)	(2)	(3)	(4)	(5)
Balassa (1965) Comp. Adv.	.041 (.017)**				-.016 (.020)
Exporter Status		-.448 (.027)***			-.438 (.027)***
Product Market Tariff			1.308 (.379)***		1.248 (.393)***
Intm. Input Tariff			-3.262 (.540)***		-3.078 (.598)***
Import Penetration				-.524 (.320)	.196 (.355)
Sector-level covariates					
Sector real exch. rate (EP^*/P)	-1.246 (.605)**	-.936 (.606)	-.935 (.626)	-.967 (.611)	-.791 (.639)
FDI Flow (USD billion)	.040 (.022)*	.047 (.021)**	.056 (.021)***	.047 (.021)**	.058 (.022)***
Herfindahl Index (sales)	-.348 (.268)	-.345 (.268)	-.796 (.282)***	-.275 (.277)	-.788 (.297)***
Plant-level covariates					
Log Employment	-.191 (.008)***	-.140 (.009)***	-.190 (.008)***	-.189 (.008)***	-.141 (.009)***
Share: Middle School or less	.948 (.107)***	.858 (.105)***	.941 (.107)***	.949 (.107)***	.851 (.105)***
Share: Some High School	.739 (.124)***	.667 (.122)***	.739 (.124)***	.740 (.124)***	.667 (.122)***
Share: White-collar occ.	-.674 (.067)***	-.613 (.067)***	-.678 (.067)***	-.671 (.067)***	-.621 (.067)***
Worker-level covariates					
Prof. or Manag'l. Occ.	-.800 (.068)***	-.806 (.068)***	-.800 (.068)***	-.799 (.068)***	-.806 (.068)***
Tech'l. or Superv. Occ.	-.601 (.064)***	-.608 (.064)***	-.595 (.064)***	-.601 (.064)***	-.603 (.064)***
Unskilled Wh. Collar Occ.	-.491 (.061)***	-.498 (.062)***	-.489 (.062)***	-.490 (.061)***	-.496 (.062)***
Skilled Bl. Collar Occ.	-.417 (.032)***	-.413 (.032)***	-.413 (.032)***	-.417 (.032)***	-.410 (.032)***
Obs.	112,978	112,978	112,978	112,978	112,978
Pseudo R^2	.036	.040	.037	.036	.041

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. Controlling for year effects. Robust standard errors in parentheses: * significance at ten, ** five, *** one percent.

workers out of manufacturing (see Table 17 in the Appendix).

In summary, Brazil's trade liberalization predicts significant changes in separation and accession rates across sectors. Penetration with foreign imports increases by 2 percentage points between 1990 and 1994, and by 6 percentage points over the full period from 1990 to 1998. The coefficient estimates thus predict a male manufacturing worker's separation odds to be 3 percentage points higher in 1994 than in 1990, and 8 percentage points higher in 1998 than in 1990. Reductions in product and input tariffs also predict significant changes in separation and accession rates, with net effects depending on the industry's dispersion of product and intermediate-input tariffs before and after reform. Between 1990 and 1998, product market tariffs drop by 19 and intermediate input tariffs by 15 percentage points on average in the estimation sample. Given the counteracting effect of the input-tariff coefficient, these drops predict an average reduction of separation rates at the mean employer. But the wide initial tariff dispersion results in a heterogenous response across sectors: the removal of trade barriers predicts a drop in the separation odds by 40 percent in footwear and 27 percent in apparel and textiles at the one extreme, but an increase in the separation odds by 9 percent in machinery manufacturing and 10 percent in transport equipment manufacturing at the other extreme. These diverse responses result in predicted labor shifts. But neither comparative-advantage sectors nor exporters exhibit the expected labor absorption. To the contrary, comparative-advantage sectors displace workers significantly more frequently than other sectors. Exporters separate from workers significantly more frequently and absorb workers significantly less frequently than nonexporters.

5.2 Reallocation durations

With evidence on labor turnover at hand, we revisit the reallocation durations from Figure 1 in the Introduction and estimate rehiring hazards for prime-age male workers after displacement from a formal-sector manufacturing job. To construct the duration sample, we record every displaced manufacturing worker between 1990 and 2001 until we observe a formal-sector rehiring.[19]

In terms of duration analysis, our setting is a multiple-record model for the duration of lacking reallocation with multiple possible successes per worker, where success is reallocation to any sector after displacement. Denote with τ the dura-

[19] If a worker who holds multiple jobs is displaced from one job but not the other, we count zero months to rehiring.

Table 10: Rehiring Hazard Estimation, 1990-2001

	Censored in 2001		Rehires within 48 months		
	(1)	(2)	(3)	(4)	(5)
Displacing sector and firm covariates					
Balassa Comp. Adv.	-.126	-.047	-.063	-.038	-.011
	(.012)***	(.008)***	(.006)***	(.007)***	(.004)***
Exporter Status	.034	.016	.059	.084	.024
	(.017)*	(.015)	(.010)***	(.011)***	(.006)***
Product Market Tariff	.780	.100	.441	.428	-.054
	(.233)***	(.304)	(.147)***	(.162)***	(.109)
Intm. Input Tariff	-2.644	-.290	-4.711	-2.865	.119
	(.500)***	(.668)	(.194)***	(.251)***	(.172)
Import Penetration	1.525	.334	3.528	2.118	.086
	(.311)***	(.207)	(.117)***	(.130)***	(.074)
Rehiring sector and firm covariates					
Balassa Comp. Adv.				-.024	-.003
				(.007)***	(.003)
Exporter Status				-.084	-.061
				(.010)***	(.006)***
Product Market Tariff				-.344	.200
				(.146)**	(.108)*
Intm. Input Tariff				-1.750	-.426
				(.225)***	(.159)***
Import Penetration				1.933	.114
				(.121)***	(.068)*
Year effects		yes			yes
Obs.	58,623	58,623	19,361	19,361	19,361

Source: *RAIS* 1990-2001. Male workers nationwide (1% random sample), 25 to 64 years old (in highest paying job if many), displaced from formal-sector manufacturing job between 1990 and 1997. Maximum-likelihood estimation of gamma distributed accelerated "failure time" (success) model for formal job reallocation in any sector before December 31, 2001 (censored sample) or into manufacturing job within 48 months (rehires sample). Sector information at subsector IBGE level. Further regressors (not reported): Sector, plant and worker covariates. Robust standard errors in parentheses: * significance at ten, ** five, *** one percent.

tion of a worker in the pool to be reallocated: τ is a continuous random variable with the cumulative distribution function $F(t) = Pr(\tau \leq t)$. The probability of a worker remaining unallocated at time t is $S(t) = Pr(\tau > t)$.[20] We specify an accelerated success time model with

$$\ln t_i^A = \mathbf{z}_{S(J(i))}\beta_S + \mathbf{y}_{J(i)}\beta_J + \mathbf{x}_i\beta_x + \alpha_T + \eta_i, \tag{3}$$

where $\mathbf{z}_{S(J(i))}$ is a vector of sector-level covariates of the worker's displacing sector $S(J(i))$; $\mathbf{y}_{J(i)}$ is a vector of plant-level covariates of worker i's displacing plant $J(i)$; $\mathbf{x}_i$ is a vector of covariates that are specific to the worker or displacing job; β_S, β_J, β_x are coefficient vectors; α_T is a year effect for $T - 1 \leq t_i < T$ and η_i is a worker-specific error with a three-parameter generalized gamma distribution. The generalized gamma distribution includes the Weibull, exponential, and log normal distributions as special cases.[21] The last observed accession in our version of *RAIS* is in December 2001, so we censor the sample at that date and estimate the duration model.

In an alternative specification, we condition on rehiring sector characteristics after restricting the sample to workers with successful reallocations to a manufacturing job within 48 months:

$$\ln t_i^B = \mathbf{z}_{R(H(i))}\beta_R + \mathbf{y}_{H(i)}\beta_y + \ln t_i^A, \tag{4}$$

where $\mathbf{z}_{R(J(i))}$ is a vector of sector-level covariates of the worker's rehiring sector $R(H(i))$; $\mathbf{y}_{H(i)}$ is a vector of plant-level covariates of worker i's rehiring plant $H(i)$; β_R and β_H are coefficient vectors and $\ln t_i^A$ is as specified in (3). This specification removes potential omitted variable bias stemming from the fact that the displacing manufacturing sector is frequently also the rehiring sector (Table 4).

Table 10 shows results from maximum-likelihood estimation. Columns 1 and 2 present estimates for the censored sample under specification (3). Censored workers may or may not find re-employment after December 2001. Workers displaced from a comparative-advantage sector have a significantly lower chance of being

[20] The reallocation hazard is thus $\lambda(t) = f(t)/S(t)$, where $f(t)$ is the density corresponding to $F(t)$, so that $S(t) = \exp\{-\Lambda(t)\}$ for $\Lambda(t) \equiv \int_0^t \lambda(u)du$.

[21] Tests overwhelmingly reject the proportionality assumption of the Cox proportional hazards model in our data. The two ancillary generalized-gamma parameters are statistically significant. We also calculate Akaike's information criterion for five distributional assumptions (exponential, Weibull, log normal, log logistic, and generalized gamma) under the accelerated success time model. We find the generalized gamma distribution to receive most empirical support in regressions with year effects, and to rank second only to the log logistic model in regressions without year effects.

reallocated. Recall that separation regressions show comparative-advantage industries to displace relatively more workers than the average sector. Together with the fact that most workers are reallocated within their sector, this suggests that workers displaced from a comparative-advantage industry have a worse reallocation chance. Duration estimation reinforces this implication. Year effects (column 2) annihilate statistical significance at the five-percent levels for all other trade variables.

Column 3 repeats column 1 estimation for displaced workers with a successful reallocation within 48 months and shows that the sample restriction hardly alters significance and sign patterns compared to the censored sample. Neither does inclusion of rehiring sector and firm characteristics in column 4 change signs or significance compared to column 3. For tariffs and import penetration, however, correlations continue to be due to time variation only and not to any sectoral component (column 5). Similar to the unrestricted sample, displacement from a comparative-advantage sector is associated with a significantly lower reallocation hazard. In contrast, displacement from an exporter predicts a significantly higher chance of reallocation for workers who are ultimately rehired. Yet, workers with a successful reallocation are significantly less likely to be rehired by an exporter. This finding is similar to our prior evidence that accession rates are significantly lower at exporters. Elevated product tariffs in the rehiring sector are associated with a better rehiring chance (at the ten-percent significance level), similar to the sign from accession regression, whereas higher intermediate-input tariffs predict a significantly lower reallocation chance, also similar to accession regression. Strong import penetration in the rehiring sector is associated with a positive rehiring chance but the estimate is not significant at the five-percent level, similar to accession regression. Overall, the similarity of duration results to separation and accession regressions underscores empirically that our focus on separations and accessions, which burden and unburden the pool of workers to be reallocated, is adequate.

6 Concomitant Economic Changes and Reforms

Brazil's trade liberalization predicts significant changes in separation and accession rates across sectors. But neither comparative-advantage sectors nor exporters exhibit the expected labor absorption; they separate from their workers significantly more frequently than other sectors and firms, and exporters hire significantly less frequently. Estimates of work status transitions and realloca-

Table 11: Conditional Logit Estimation with Labor Productivity

	Separations			Accessions		
	Full smpl.	*PIA* smpl.		Full smpl.	*PIA* smpl.	
	(1)	(2)	(3)	(4)	(5)	(6)
Balassa Comp. Adv.	.169	-.006	-.009	-.016	-.017	-.013
	(.024)***	(.074)	(.075)	(.020)	(.060)	(.060)
Exporter Status	.283	.030	.031	-.438	-.293	-.295
	(.028)***	(.076)	(.076)	(.027)***	(.075)***	(.075)***
Prod. Market Tariff	-.715	1.240	1.228	1.248	-.317	-.269
	(.426)*	(.987)	(.987)	(.393)***	(.955)	(.958)
Intm. Input Tariff	2.901	.391	.448	-3.078	-1.285	-1.495
	(.678)***	(1.607)	(1.609)	(.598)***	(1.374)	(1.376)
Import Penetration	1.252	1.324	1.280	.196	.467	.591
	(.388)***	(1.000)	(1.002)	(.355)	(1.098)	(1.098)
Log Labor Prod.			.008			-.020
			(.008)			(.007)***
Obs.	145,417	40,337	40,337	112,978	20,185	20,185
Pseudo R^2	.150	.335	.335	.041	.089	.090

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

tion durations confirm these predictions. But empirical concerns remain. We address the role of firm-level labor productivity, the potential simultaneity of trade policies and exporting status, the relevance of concomitant reforms and the importance of unobserved worker heterogeneity.

Firm-level labor productivity. Exporters are more productive than nonexporters (Table 3). To compare the relative importance of a firm's exporter status and labor productivity for separations and accessions, we include a measure of firm-level labor productivity in specification (2). For this purpose, we use the subsample of *RAIS* firms that are surveyed in *PIA*, for which firm-level labor productivity is inferrable. Table 12 redisplays conditional logit estimates for separations and accessions on the full sample in columns 1 and 4 (from Tables 8 and 9, column 6). The table compares those prior estimates to estimates on the combined *PIA-RAIS* subsample without (columns 2 and 5) and with log labor

Table 12: Linear and Instrumental-Variable Estimation

	Separations			Accessions		
		OLS-FE			OLS-FE	
	Cdl. logit		IV	Cdl. logit		IV
	(1)	(2)	(3)	(4)	(5)	(6)
Balassa Comp. Adv.	.169	.017	.023	-.016	.002	-.002
	(.024)***	(.002)***	(.003)***	(.020)	(.002)	(.003)
Exporter Status	.283	.038	.515	-.438	-.049	-.499
	(.028)***	(.003)***	(.096)***	(.027)***	(.003)***	(.091)***
Product Market Tariff	-.715	-.100	-.033	1.248	.124	.114
	(.426)*	(.035)***	(.081)	(.393)***	(.032)***	(.073)
Intm. Input Tariff	2.901	.344	.164	-3.078	-.309	-.229
	(.678)***	(.054)***	(.141)	(.598)***	(.049)***	(.132)*
Import Penetration	1.252	.052	.003	.196	.088	.264
	(.388)***	(.034)	(.077)	(.355)	(.031)***	(.071)***
Obs.	145,417	293,358	293,358	112,978	293,118	293,118

Source: *RAIS* 1990-98 (1% random sample), male workers nationwide, 25 to 64 years old, with manufacturing job. Separations and accessions exclude transfers, deaths, and retirements. Reference observations are employments with no reported separation or accession in a given year. Sector information at subsector IBGE level. Estimates in column 1 and 4 repeat column 6 in Tables 8 and 9. Further regressors (not reported): Year indicators, sector, plant and worker covariates. Robust standard errors in parentheses: * significance at ten, ** five, *** one percent.

productivity as a regressor (columns 3 and 6). The export-status coefficient loses statistical significance in the reduced separation subsample of *PIA* manufacturers but does not change sign. Exporters exhibit significantly fewer accessions at the one-percent level; this reinforces our prior finding. Trade-variables, including Balassa comparative advantage, are not significant predictors of separations and accessions in the reduced subsample. Higher labor productivity itself predicts significantly fewer accessions. This is consistent with the descriptive evidence (Table 2) that faster labor productivity growth at manufacturing firms correlates with slower-than-average workforce growth. So, the inclusion of log labor productivity in a smaller random sample of manufacturers significantly reinforces some results and overturns none of our findings.

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 does give rise to a simultaneity concern. By design, initially highly protected sectors face the largest product tariff declines. Similarly, market penetration with foreign inputs possibly responds to Brazilian labor-market conditions. We therefore predict tariffs and market penetration rates at the sector level with instrumental variables. At the firm-level, employers decide exporting status and labor turnover simultaneously. We therefore also predict export status with instrumental variables.

To construct instruments for export demand, we consider seven broad destination regions of Brazil's exports and calculate the destination-region imports from other source-countries than Brazil. These foreign demand proxies vary by sector and year. In addition, we employ the nominal U.S. dollar exchange rate and sector-level producer-price indices in the U.S. and the EU as instruments. Nominal exchange rate movements are largely unpredictable, and foreign producer prices in industrialized economies are arguably exogenous to Brazil. To check for potential sign reversals and assess the magnitude of possible simultaneity bias, we resort to linear fixed-effects regressions of separation and accession indicators on the same predictors as in the preceding section:[22]

$$\sigma_{i,t} = \hat{\mathbf{z}}_{S(J(i)),t}\beta_z + \hat{\mathbf{y}}_{J(i),t}\beta_y + \mathbf{x}_{i,t}\beta_x + \alpha_i + \alpha_t, \tag{5}$$

where $\sigma_{i,t} \in \{0, 1\}$ denotes the binary outcome (accession or not, separation or not) for worker i at time t and regressor and coefficient vectors are as in (2). There is an unobserved error 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 (5).

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 18 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 14 and 14,000. Almost invariably, the instruments are statistically significant predictors at the one-percent level.[23] We highlight a few coefficient estimates.

[22] 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).

[23] We also experiment with labor productivity in the initial year 1990 as a candidate firm-level

Expectedly, higher producer prices in the U.S. and Europe, as well as a weaker Brazilian currency, predict significantly more frequent exporting status. Employment frequencies at exporters are predicted to be higher in sectors with weaker comparative advantage, as documented in data Section 2 before, because there is a larger number of small-volume exporters in the low-advantage sectors.

Table 12 redisplays conditional logit estimates for separations and accessions in columns 1 and 4 (from Tables 8 and 9, column 6). We compare those estimates to linear worker-fixed effects regressions without (columns 2 and 5) and with instrumental-variable predictions (columns 3 and 6). The estimation samples for the linear worker-fixed effects models are substantially larger because workers with no transition remain in the sample; their fixed effect is identified from time variation at the same employer. When instrumenting, there is not a single sign reversal in the potentially simultaneity-afflicted coefficients—export status, tariffs, and import penetration (comparing columns 2 and 3, and columns 5 and 6). Instrumentation overwhelmingly reinforces at the one-percent significance level that comparative-advantage sectors and exporters exhibit more separations, and exporters exhibit fewer accessions. Several coefficients on tariffs and import penetration lose significance at common levels under instrumental-variable fixed-effects regressions (columns 3 and 6) but never exhibit a sign reversal. So, instrumentation in a linear probability model corroborates our main explanation for lacking labor reallocation: firms in comparative-advantage sectors and exporters separate from their workers significantly more frequently than the average employer, and exporters hire significantly less frequently.

Economic change and reforms. The Brazilian economy undergoes a series of concomitant economic transformations during the sample period, including skill-biased technological change, the intensified outsourcing of service jobs, surging foreign direct investment inflows and policy shifts such as macroeconomic stabilization, capital-account liberalization, and privatization. Labor-market institutions were altered preceding trade reform. All accession, separation and reallocation-duration regressions so far control for sector and year covariates including sectoral real exchange rates, Herfindahl sales concentration indices, foreign direct investment inflows, and year fixed effects. We turn to economic changes and policies that may affect estimates at the level of the plant, job, worker or employer-employee match in specification (2).

instrument in the subsample of *PIA* firms but over-identification tests reject its validity when added.

Table 13: FURTHER CONDITIONAL LOGIT SPECIFICATIONS

	Cdl. logit	Primary school	High school	College educ.	Sector FE	Privat. control	Outsrc. job ind.
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
			SEPARATIONS				
Balassa Cmp. Adv.	.168	.144	.303	.225	-.095	.170	.169
	(.024)***	(.028)***	(.098)***	(.151)	(.049)	(.026)***	(.024)***
Exporter Status	.283	.296	.217	.295	.284	.283	.283
	(.028)***	(.033)***	(.091)**	(.143)**	(.028)***	(.028)***	(.029)***
Prod. Mkt. Tariff	-.710	-.503	-2.776	-1.912	-2.369	-.698	-.756
	(.426)*	(.499)	(1.355)**	(2.289)	(.476)***	(.427)	(.430)*
Intm. Input Tariff	2.893	2.479	8.373	7.705	5.166	2.887	3.024
	(.678)***	(.779)***	(2.416)***	(4.118)*	(.748)***	(.676)***	(.686)***
Imp. Penetration	1.247	.667	1.935	.814	3.217	1.255	1.260
	(.388)***	(.477)	(1.279)	(1.998)	(.638)***	(.393)***	(.391)***
addl. regressor(s)					yes	-.154	-.018
						(1.228)	(.037)
Obs.	145,418	110,846	17,627	7,493	145,418	145,418	143,546
Pseudo R^2	.150	.161	.269	.247	.151	.150	.152
			ACCESSIONS				
Balassa Cmp. Adv.	-.016	-.006	-.165	-.150	-.068	-.024	-.015
	(.020)	(.023)	(.086)*	(.118)	(.048)	(.022)	(.021)
Exporter Status	-.439	-.421	-.504	-.775	-.439	-.439	-.438
	(.027)***	(.031)***	(.093)***	(.140)***	(.027)***	(.027)***	(.027)***
Prod. Mkt. Tariff	1.248	1.336	2.533	2.281	1.820	1.118	1.187
	(.393)***	(.451)***	(1.399)*	(2.088)	(.498)***	(.412)***	(.397)***
Intm. Input Tariff	-3.078	-2.947	-8.501	-5.682	-2.952	-2.991	-3.047
	(.598)***	(.673)***	(2.292)***	(3.386)*	(.750)***	(.603)***	(.605)***
Imp. Penetration	.203	.093	.358	-.646	1.773	.132	.187
	(.355)	(.423)	(1.184)	(1.949)	(.665)***	(.363)	(.358)
addl. regressor(s)					yes	1.161	-.098
						(1.167)	(.033)***
Obs.	112,971	86,469	12,062	4,782	112,971	112,971	110,985
Pseudo R^2	.041	.043	.090	.089	.042	.041	.040

Source: *RAIS* 1990-98 (1% random sample), male workers nationwide, 25 to 64 years old, with manufacturing job. Estimates in column 1 repeat column 6 in Tables 8 and 9. Further regressors (not reported): Year indicators, sector, plant and worker covariates. Robust standard errors in parentheses: * significance at ten, ** five, *** one percent.

If skill-biased technological change systematically interacts with the effect of trade reform on labor turnover, trade reform must arguably covary with labor turnover differently for workers with different skills. We run specification (2) separately for workers with primary schooling, some high school attendance and completed college. Table 13 redisplays in column 1 the conditional logit estimates for separations and accessions on the full sample. Estimates for the education subsamples follow in columns 2 through 4. Coefficient estimates for separations and accessions are strikingly similar across the samples. No sign changes. 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 high-school and college educated worker subsamples. There is, to our surprise, no evidence that skill-biased technological change systematically interacts with the effect of trade reform on separations and accessions.

Though the constitutional labor-market reforms in 1988 precede trade reform in 1990, they might affect sectors with unobserved differences in workforce composition to varying degrees and interact with trade reform in a way that erroneously attributes labor turnover to the trade regime. We use sector-fixed effects at the subsector IBGE level to capture unobserved sectoral differences in the effect of labor institutions on unobserved separation and accession determinants. Table 13 reports estimates in column 5. Expectedly, inclusion of sector indicators turns the coefficient on comparative advantage, which is highly sector specific and largely time invariant in our data, insignificant. For the other trade regressors, however, coefficient estimates increase in absolute value (compared to column 1) and remain highly significant. This renders erroneous attribution of labor-market regulation effects to trade reform little plausible.

The privatization of state-owned businesses and the progressing outsourcing of service jobs to specialized suppliers can affect separations and accessions. If privatization and outsourcing covary with the trade regime and labor turnover in systematic ways, they potentially lead to erroneous attributions. The ownership status of a plant is observable in *RAIS* since 1995, when the federal government started to pursue privatization on a larger scale. In manufacturing industries, the employment-weighted share of private companies rises from xx percent to zz percent between 1995 and 1998. We impute a plant's ownership status in 1990-94 as the ownership status in 1995 and include the private-ownership indicator at the plant-level in regression (2). As column 6 in Table 13 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

Table 14: UNCONDITIONAL LOGIT ESTIMATION

	Separations			Accessions		
	Cdl.	Logit		Cdl.	Logit	
	logit	cdl. smpl.	full smpl.	logit	cdl. smpl.	full smpl.
	(1)	(2)	(3)	(4)	(5)	(6)
Balassa Cmp. Adv.	.169	.108	.125	-.016	.072	.184
	(.024)***	(.009)***	(.007)***	(.020)	(.009)***	(.007)***
Exporter Status	.283	.066	-.017	-.438	-.303	-.507
	(.028)***	(.015)***	(.012)	(.027)***	(.017)***	(.013)***
Prod. Mkt. Tariff	-.715	-.093	-.491	1.248	.222	-1.400
	(.426)*	(.221)	(.170)***	(.393)***	(.251)	(.198)***
Intm. Input Tariff	2.901	.975	1.375	-3.078	-.935	2.227
	(.678)***	(.344)***	(.255)***	(.598)***	(.380)**	(.301)***
Import Penetration	1.252	-.124	-.329	.196	.182	-1.004
	(.388)***	(.173)	(.140)**	(.355)	(.203)	(.165)***
Obs.	145,417	145,417	293,358	112,978	112,978	293,118
Pseudo R^2	.150	.033	.050	.041	.023	.078

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 columns 1 and 4 repeat column 6 in Tables 8 and 9. Further regressors (not reported): Year indicators, sector, plant and worker covariates. Robust standard errors in parentheses: * significance at ten, ** five, *** one percent.

susceptibility of a job to outsourcing (*tercerização*) if it is a service occupation at the CBO three-digit level that can be performed in-house or be provided by a specialized subcontractor. Including the job-level indicator in regression (2) results in no statistically significant coefficient change (column 7). Jobs susceptible to outsourcing exhibit a statistically significant reduction 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.

Unobserved worker heterogeneity. Worker heterogeneity is an important predictive component of separations and accessions. A comparison between conditional logit estimation and logit estimation without worker-fixed effects in Table 14 shows that logit is sensitive to omission of the worker-fixed effect. Import penetration loses statistical significance (and reverts sign) in separation regression (comparing column 1 to 2). Comparative advantage reverts sign and turns statistically significant in accession regression, while product-market tariffs lose significance (comparing column 4 to 5). Recall from our comparison between conditional logit estimation and linear worker-fixed effects estimation that there is no single significant sign reversal between those specifications although the sample widens because workers without transition remain in the linear-regression sample (comparing column 2 to 1 and 4 to 3 in Table 12). Logit estimation, in contrast, is sensitive to the widening sample: export status loses significance (and changes sign) in separation regression (comparing column 2 to 3), and so do product tariffs and import penetration in accession regression as the sample widens (column 5 to 6). Input tariffs even change sign from significantly negative to significantly positive in logit accession estimation as the sample widens. These findings are consistent with the hypothesis that the termination and formation of matches between individual workers and employers is not random, even after conditioning on a comprehensive set of observable worker and employer characteristics.

7 Labor Market Evidence and Trade Theory

Our finding that neither comparative-advantage sectors nor exporters absorb displaced workers after trade reform stands in stark contrast with classic trade theory (Ricardo, Heckscher-Ohlin-Samuelson) and recent firm-level trade models (Bernard et al. 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.[24] Our research documents, however, that the main concern is not lacking employment reduction; worker separations significantly increase with import penetration. It is rather the paucity of formal-sector re-accessions after separations that characterizes the failed reallocation process.

Aspects of Brazil's experience might be perceived as consistent with predictions of recent trade models that make factor-market institutions a source of

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

comparative advantage and find that countries with less rigid factor markets tend to specialize in industries with high factor turnover (Saint Paul 1997, Davidson et al. 1999, Cunat and Melitz 2006). Brazil's comparative-advantage sectors indeed exhibit more labor turnover: significantly higher worker separation rates and, unconditionally, higher worker accession rates. The lacking net expansion of comparative advantage sectors, however, is not compatible with that explanation. Moreover, comparing World Bank indices of labor-market rigidity for Brazil to weighted averages of Brazil's trading partners shows that Brazil's labor market is considerably more rigid.[25] So, those theories would predict Brazil to specialize in industries with low labor turnover, contrary to our evidence.

Bernard et al. (2006b) 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.[26]

An empirical investigation of the Bernard et al. model's labor-market predictions calls for the inclusion of higher-order interactions between trade reform, comparative advantage and exporting status. Table 15 compares our previous separation and accessions estimates in columns 1 and 4 (from Tables 8 and 9, 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 in the Bernard et al. model that

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

[26]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.

Table 15: CONDITIONAL LOGIT ESTIMATION WITH INTERACTIONS

	Separations			Accessions		
	(1)	(2)	(3)	(4)	(5)	(6)
Balassa Comp. Adv.	.169	.139	.134	-.016	-.059	-.125
	(.024)***	(.036)***	(.043)***	(.020)	(.032)*	(.038)***
Comp. Adv. $\times$ Prd. Trff.		.197	.264		.290	.598
		(.200)	(.238)		(.162)*	(.203)***
Exporter Status	.283	.481	.476	-.438	-.360	-.562
	(.028)***	(.048)***	(.081)***	(.027)***	(.045)***	(.077)***
Exporter $\times$ Prd. Trff.		-1.070	-.942		-.424	.348
		(.213)***	(.362)***		(.195)**	(.323)
Comp. Adv. $\times$ Exporter			.012			.154
			(.051)			(.047)***
Comp. Adv. $\times$ Exp. $\times$ Prd. Trff.			-.148			-.675
			(.291)			(.250)***
Product Market Tariff	-.715	-.427	-.506	1.248	.966	.543
	(.426)*	(.532)	(.549)	(.393)***	(.474)**	(.504)
Intm. Input Tariff	2.901	3.253	3.302	-3.078	-2.490	-2.301
	(.678)***	(.768)***	(.767)***	(.598)***	(.672)***	(.682)***
Import Penetration	1.252	1.091	1.085	.196	.033	-.003
	(.388)***	(.393)***	(.393)***	(.355)	(.364)	(.364)
Obs.	145,417	145,417	145,417	112,978	112,978	112,978
Pseudo R^2	.150	.151	.151	.041	.041	.041

Source: *RAIS* 1990-98 (1% random sample), male workers nationwide, 25 to 64 years old, with manufacturing job. Separations and accessions exclude transfers, deaths, and retirements. Reference observations are employments with no reported separation or accession in a given year. Sector information at subsector IBGE level. Further regressors (not reported): Year indicators, sector, plant and worker covariates. Columns 1 and 4 repeat column 6 of Tables 8 and 9. Robust standard errors in parentheses: * significance at ten, ** five, *** one percent.

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. model: in comparative-advantage industries, existing jobs are destroyed less frequently at (high-productivity) exporters.

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

8 Conclusions

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

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

Appendix

A Linked employer-employee data

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

Observation screening. In *RAIS*, workers are identified by an individual-specific PIS (*Programa de Integração Social)* ID number that is similar to a social security number in the U.S. (but PIS IDs are not used for identification purposes other than the administration of the wage supplement program *Abono Salarial*). A given plant may report the same PIS ID 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 IDs to be recorded incorrectly or repeatedly. To handle these issues, we screen *RAIS* in two steps. (1) Observations with PIS IDs shorter than 11 digits are removed. These may correspond to informal (illegal) workers or measurement error from faulty bookkeeping. (2) For several separation statistics, we remove multiple jobs from the sample if a worker's duplicate jobs have identical accession and separation dates at the same plant. For a worker with such multiple employments, we only keep the observation with the highest average monthly wage level (in cases of wage level ties, we drop duplicate observations randomly).

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

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

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

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

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

Occupation indicators derive from the 3-digit CBO classification codes in our nationwide *RAIS* data base, and are reclassified to conform to the ISCO-88 categories.[27] We map ISCO-88 categories to *RAIS* occupations as follows:

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

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

B Manufacturing firm data

For robustness checks in Table 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 log TFP to log labor productivity by adding the production-coefficient weighted effects of capital accumulation and intermediate input use.

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

C Additional sector data

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

We use Ramos and Zonenschain (2000) national accounting data to calculate the effective rate of market penetration with foreign imports. Arguably, domestic firms find the absorption market corresponding to $A_{i,t} \equiv Y_{i,t} - (X_{i,t} - M_{i,t})$ the relevant domestic environment in which they compete. We define the effective rate of market penetration as $M_{i,t}/A_{i,t}$. Foreign direct investment (FDI) flow 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 low value means an appreciated 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 indices are not available in general.

Table 16: SUBSECTOR IBGE AND *Nível 50* COMPARISON

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

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

Table 17: Year Effects in Conditional Logit Estimates

	Separations			Accessions		
	(1)	(2)	(3)	(4)	(5)	(6)
Year 1990	-2.064	-2.128	-2.133	1.035	.966	.953
	(.136)***	(.145)***	(.145)***	(.126)***	(.131)***	(.131)***
Year 1991	-1.326	-1.357	-1.357	1.261	1.226	1.217
	(.067)***	(.070)***	(.070)***	(.062)***	(.064)***	(.064)***
Year 1992	-.969	-.979	-.978	1.099	1.087	1.082
	(.110)***	(.110)***	(.110)***	(.109)***	(.109)***	(.109)***
Year 1993	-.859	-.860	-.859	1.121	1.124	1.122
	(.067)***	(.067)***	(.067)***	(.067)***	(.067)***	(.067)***
Year 1994	-.863	-.858	-.858	.972	.985	.989
	(.047)***	(.047)***	(.047)***	(.047)***	(.047)***	(.047)***
Year 1995	-.445	-.432	-.433	.699	.723	.731
	(.085)***	(.086)***	(.086)***	(.086)***	(.087)***	(.087)***
Year 1996	-.377	-.368	-.368	.686	.701	.706
	(.050)***	(.050)***	(.050)***	(.052)***	(.052)***	(.052)***
Year 1997	-.204	-.195	-.194	.489	.502	.506
	(.039)***	(.040)***	(.040)***	(.041)***	(.041)***	(.041)***
Trade-related covariates						
2nd order interactions		yes			yes	
3rd order interactions		yes	yes		yes	yes

Source: *RAIS* 1990-98 (1% random sample), male workers nationwide, 25 to 64 years old, with manufacturing job. Separations and accessions exclude transfers, deaths, and retirements. Reference observations are employments with no reported separation or accession in a given year. Year effects from conditional logit estimation: column 1 completes column 6 of Table 8, columns 2 and 3 complete columns 2 and 3 of Table 15, column 4 completes column 6 of Table 9, columns 5 and 6 complete columns 5 and 6 of Table 15. Other regressors (not reported): Trade-related, sector (subsector IBGE level), plant and worker covariates. Robust standard errors in parentheses: * significance at ten, ** five, *** one percent.

Table 18: 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	3.526 (.789)***	-2.279 (.097)***	-.004 (.053)	3.822 (.975)***	-2.123 (.111)***	.386 (.065)***
World imports CEE	43.524 (4.341)***	-33.883 (.534)***	-16.622 (.293)***	38.807 (5.550)***	-26.907 (.635)***	-17.060 (.370)***
World imports LAC	-4.757 (1.035)***	14.263 (.127)***	4.757 (.070)***	-2.027 (1.319)	14.039 (.151)***	4.864 (.088)***
World imports NAM	-2.357 (.525)***	-.651 (.065)***	-1.670 (.035)***	-2.438 (.662)***	.378 (.076)***	-1.992 (.044)***
World imports ODV	-2.101 (.763)***	-5.736 (.094)***	.311 (.052)***	-1.353 (.977)	-5.274 (.112)***	-.138 (.065)**
World imports OIN	4.182 (.957)***	-9.099 (.118)***	-5.675 (.065)***	4.015 (1.181)***	-10.353 (.135)***	-5.341 (.079)***
World imports WEU	13.954 (.461)***	2.159 (.057)***	1.949 (.031)***	14.438 (.564)***	1.468 (.065)***	2.096 (.038)***
PPI Idx. EU	.706 (.115)***	-.928 (.014)***	.112 (.008)***	.977 (.144)***	-.940 (.016)***	.052 (.010)***
PPI Idx. NAM	.412 (.106)***	.850 (.013)***	-.120 (.007)***	.474 (.138)***	.802 (.016)***	-.200 (.009)***
USD Exch. Rate	.106 (.025)***	-.211 (.003)***	.011 (.002)***	.081 (.032)**	-.252 (.004)***	-.014 (.002)***
Exogenous covariates						
Balassa Comp. Adv.	-.020 (.003)***	-.026 (.0003)***	-.022 (.0002)***	-.024 (.003)***	-.027 (.0004)***	-.022 (.0002)***
FDI Flow (USD billion)	.002 (.003)	.014 (.0004)***	.004 (.0002)***	.0001 (.004)	.014 (.0004)***	.005 (.0003)***
Herfindahl Index (sales)	.328 (.044)***	.048 (.005)***	.053 (.003)***	.252 (.054)***	-.026 (.006)***	.098 (.004)***
Log Employment	.052 (.002)***	.003 (.0002)***	-.0009 (.0001)***	.050 (.002)***	.003 (.0002)***	-.0007 (.0001)***
Share: Middle Sch. or less	-.171 (.016)***	.008 (.002)***	-.007 (.001)***	-.184 (.017)***	.007 (.002)***	-.009 (.001)***
Share: Some High School	-.062 (.019)***	-.002 (.002)	.003 (.001)**	-.092 (.021)***	-.005 (.002)**	.002 (.001)
Share: White-collar occ.	.060 (.010)***	.006 (.001)***	-.002 (.0007)**	.056 (.012)***	.004 (.001)***	-.002 (.0008)**
F statistic (IV)	13.531	14,340.84	475.562	23.906	12,721.4	310.205

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 8 for separations, Table 9 for accessions), controlling for year effects. Annual sector-weighted world imports, coefficients rescaled to imports in USD trillion. Robust standard errors in parentheses: * significance at ten, ** five, *** one percent.

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