Trade Reform, Employment Allocation and Worker Flows*

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*This summary report synthesizes previously circulated research in Menezes-Filho, Muendler and Ramey (2008), Muendler (2008), Muendler (2004) and Menezes-Filho and Muendler (2007), partly joint with Naércio Aquino Menezes-Filho and Garey Ramey. The report was completed while the author was visiting Princeton University. Financial support from the World Bank is gratefully acknowledged.

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1 Introduction

This report summarizes insights from research into Brazil's labor-market adjustment following the country's large-scale trade reform in the early 1990s, synthesizing findings from Menezes-Filho et al. (2008), Muendler (2008), Muendler (2004) and Menezes-Filho and Muendler (2007).

Two salient workforce changeovers become evident from a labor-demand decomposition based on the Katz and Murphy (1992) framework. Within the traded-goods sector, there is a marked occupation downgrading and a simultaneous education upgrading by which employers fill expanding low-skill intensive occupations with increasingly educated jobholders. Between sectors, there is a labor demand shift towards the least and the most skilled, which can be traced back to relatively weaker declines of traded-goods industries that intensely use low-skilled labor and to relatively stronger expansions of nontraded-output industries that intensely use high-skilled labor. Interestingly, and in a certain contrast to the experience of other Latin American economies, these observations are broadly consistent with predictions of Heckscher-Ohlin trade theory for a low-skill abundant economy.

The conventional decomposition leaves unaddressed, however, how workforce changeovers come about. For this purpose, actual worker flows need to be observed within and across employers. Rich linked employer-employee data help address this question and show that workforce changeovers are neither achieved through worker reassignments to new tasks within employers nor are they brought about by reallocations across employers and traded-goods industries. Instead, trade-exposed industries shrink their workforces by dismissing less-schooled workers more frequently than more-schooled workers especially in skill-intensive occupations, while most displaced workers shift to nontraded-output industries or out of recorded employment. Brazil's trade liberalization triggers worker displacements particularly from protected industries, as trade theory predicts and welcomes. But neither comparative-advantage industries nor exporters absorb trade-displaced workers for years. In fact, comparative-advantage industries and exporters displace significantly more workers and hire fewer workers than the average employer, and resource reallocation appears to remain incomplete for years. These patterns pose a challenge to classic trade theory, including the Heckscher-Ohlin framework, as well as modern trade theories with heterogeneous firms in the absence of endogenous productivity change.

To investigate explanations for the reverse labor flows away from comparative-advantage sectors and away from exporters more closely, I combine the linked employer-employee data with data from a Brazilian manufacturing survey. The combined data set shows that labor is flowing away from comparative-advantage sectors and away from exporters because their labor productivity increases faster than their production so that output shifts to more productive firms, as has been widely documented, while labor does not, contrary to often hypothesized resource flows. The most plausible explanation seems to be that trade triggers faster productivity growth at exporters and in comparative-advantage industries because for these firms and industries larger market potential offers stronger incentives to improve efficiency. If productivity increases faster than production, then output shifts to more productive firms but labor does not.

The labor-market evidence for Brazil also offers a novel explanation why pro-competitive reforms can be associated with strong efficiency gains at the employer level but not in the aggregate, where idle resources result. Conservatively measured, the foregone wage bill from the increase in reallocation durations and failures after 1990 amounts to between one and three percent of GDP. The increase in joblessness is not solely due to trade integration. But regression analysis at the worker level, and a series of robustness checks to rule out alternative explanations, document that trade variables predict a large part of the fluctuation in displaced labor.

This report focuses on Brazil's labor reallocation for prime-age male workers, 25 to 64 years old, in the formal sector anywhere nationwide. The restriction to prime-age workers is meant to reduce the sample to workers after their first labor-force entry, highlighting the re-allocation of active labor resources. Prime-age males workers are known to exhibit low wage elasticities of labor supply so that the presented results are possibly little affected by labor supply changes. A recent revision to Menezes-Filho and Muendler (2007) shows that results are similar for samples that include both genders and all age groups.

The empirical literature on trade and resource reallocation has taken mainly three approaches. First, industry-level studies use measures of job creation, destruction, and churning (excess turnover beyond net change), as well as informality. Haltiwanger, Kugler, Kugler, Micco and Pagés (2004) show for a panel of six Latin American countries, for instance, that tariff reductions are associated with heightened within-sector churning and net employment reductions at the sector level.[1] Beyond those studies, research with linked employer-employee as discussed in this report documents the direction of factor flows between types of employers, and identifies the incidence of idle resources in the process. In contrast to the United States, where industries with faster productivity growth exhibit higher net employment growth (Davis et al. 1996), more productive employers reduce employment in Brazil during the 1990s. Using sector data, Goldberg and Pavcnik (2003) report no statistically significant relation between informal work and trade in Brazil, whereas household survey data suggest that tariff reductions are related to more transitions out of formal work, especially into self-employment and withdrawals from the labor force (Menezes-Filho and Muendler 2007). The present research report, however, focuses on the formal sector.

Second, employer-level studies show that trade reforms are associated with product-market reallocation towards more efficient producers (for a survey see Tybout 2003). But employer-level studies typically report no detectable relationship between trade and employment.[2] As to evidence from Brazil discussed in this research report, trade variables are

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

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

not statistically significant predictors of employment changes at the employer-level either (Muendler 2008). But worker-level regressions on the same data uncover that additional imports trigger significantly more worker displacements, while there are lasting worker flows away from productive high-output employers. This suggests that unobserved workforce heterogeneity hampers regressions at more aggregate levels, even the employer level, and calls for the use of worker panel data.

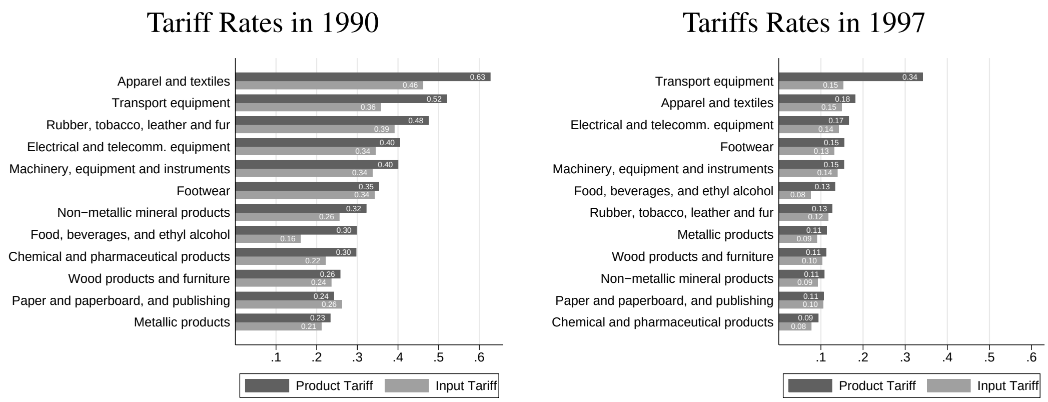
Third, a worker-level literature studies the experience of displaced workers across sectors and worker groups. Kruse (1988) and Kletzer (2001) compare displaced workers between U.S. industries and find that employment histories are largely explained by differences in workforce characteristics across sectors and vary little by a sector's trade exposure.[3] Time variation in our data, by contrast, identifies a salient impact of Brazil's trade opening on labor turnover. Beyond displaced-worker survey data, our linked employer-employee records allow us to quantify directions of worker flows across employers for many years and show that the economic burden of joblessness is substantial. As this report documents, the joblessness is partly trade induced.

The remainder of this report is organized as follows. Section 2 briefly summarizes Brazil's trade reform and compares the country's labor-market characteristics to other economies. Section 3 introduces the data (with most details relegated to the Appendix). Section 4 presents labor demand changes over the sample period 1986-2001 and discerns between-sector and within-sector changes using a Katz and Murphy (1992) labor demand decomposition. This exercise documents the workforce changeover within sectors along educational and occupational dimensions but leaves unaddressed which worker flows are associated with the observed employment changes. Section 5 investigates how much of the documented workforce changeover is brought about by task reassignments within firms, worker reallocations across firms and industries, and by worker separations without formal-sector reallocations. Section 6 uses a regression design, controlling for worker heterogeneity in turnover, to identify what share of the reallocation flows during the 1990s is predicted by Brazil's heightened trade exposure. Section 7 offers as a main explanation for the observed reverse labor flows away from comparative-advantage sectors and away from exporters that labor productivity changes endogenously in response to trade reform, and presents according statistics. Section 8 discusses potential implications of the findings for labor-market adjustment costs. Section 9 concludes.

2 Brazil and its Trade Reform

Since the late 1980s, Brazil's federal government initiated a series of economic reforms that by around 1997 resulted in a considerably more open economy to foreign goods and investments, a stable macroeconomy, and a somewhat smaller role of the state in the economy. In 1988, after decades of import substitution and industry protection, the Brazilian federal government under president Sarney initiated an internal planning process for trade reform

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



Sources: Muendler (2008). *Ad valorem* product tariffs at *Nível 80* from Kume et al. (2003).
Note: Intermediate input tariffs are re-weighted product-market tariffs using national input-output matrices at *Nível 80* (from *IBGE*). Product-market and intermediate-input tariffs are transformed to the subsector *IBGE* level using unweighted means over the *Nível 80* classifications.

Figure 2.1: **Product-market and intermediate-input tariffs 1990 and 1997**

and started to reduce *ad valorem* tariffs but, lacking public support, took little legislative initiative to remove binding non-tariff barriers so that nominal tariff reductions had little effect (Kume, Piani and Souza 2003). In 1990, the Collor administration launched a large-scale trade reform that involved both the removal of non-tariff barriers and the adoption of a new tariff structure with lower levels and smaller cross-sectoral dispersion. As a surprise to most observers at the time, Collor abolished all non-tariff barriers by presidential decree on his first day in office. Implementation of these policies was largely completed by 1993.

Figure 2.1 depicts Brazil's product-market and intermediate-input tariff schedules in 1990 and 1997 for the twelve manufacturing industries at the subsector *IBGE* level. Intermediate input tariff levels are calculated as re-weighted product tariffs using the economy-wide input-output matrix. Both the level and the dispersion of tariffs drop remarkably between 1990 and 1997. While *ad valorem* product tariffs range from 21 (metallic products) to 63 percent (apparel and textiles) in 1990, they drop to a range from 9 percent (chemicals) to 34 percent (transport equipment) in 1997. Except for paper and publishing in 1990, sectors at the subsector *IBGE* level receive effective protection in both years, with mean product tariffs exceeding mean intermediate-input tariffs. By 1997, however, the relatively homogeneous tariff structure results in a small rate of effective protections for most industries—with the notable exception of transport equipment.

Brazil underwent additional reforms over the sample period. In 1994, during the Franco administration and under the watch of then finance minister Cardoso, drastic anti-inflation measures succeeded for the first time in decades. A privatization program for public utilities was started in 1991 and accelerated in the mid 1990s, while Brazil simultaneously liberalized capital-account restrictions. These measures were accompanied by a surge in foreign direct investment inflows in the mid 1990s. The pro-competitive reforms during the 1990s, mostly targeted at product markets, had been preceded by changes to Brazil's labor-market institutions in 1988.

Table 2.1: LABOR MARKET RIGIDITY COMPARISONS

	Rigidity and Difficulty Indices				
	Hiring difficulty	Hours rigidity	Firing difficulty	Employment rigidity	Firing costs[a]
	(1)	(2)	(3)	(4)	(5)
Brazil	67.0	80.0	70.0	72.0	165.0
Trade partners					
weighted by trade volume[b]					
1990	25.2	42.0	22.7	29.9	43.3
1997	28.1	45.3	24.4	32.4	47.6
weighted by source-country imports					
1990	23.2	42.9	21.7	29.1	46.8
1997	27.2	44.3	23.6	31.6	46.0
weighted by destination-country exports					
1990	26.4	41.5	23.4	30.3	41.2
1997	29.1	46.4	25.2	33.4	49.5

[a]In weekly wage equivalents.
[b]Country sum of exports from and imports to Brazil.

Source: Botero, Djankov, La Porta, Lopez de Silanes and Shleifer (2004) labor market rigidity measures.
Note: A higher index and a higher rank indicate a more rigid labor market. Trade partner averages weighted by *WTF* (NBER) bilateral trade data for 1990 and 1997.

Brazil's constitution of 1988 introduced a series of labor-market reforms that aimed to increase workers' benefits and the right to organize, thus raising labor costs. Most important, firing costs increased substantially.[4] Given their constitutional status, these labor-market institutions remained unaltered throughout the 1990s, the period of chief interest for this report. Table 2.1 compares World Bank indices of labor-market rigidity for Brazil to Brazil's mean trading partner and shows that Brazil's labor market is considerably more rigid than its trade partners' labor markets are. For the World Bank's four rigidity and difficulty indices (hiring difficulty, hours rigidity, firing difficulty, employment rigidity) and its firing-cost measure, Brazil exhibits mean values between 67 and 165, whereas the mean values for Brazil's trading partners vary between 20 and 49 for three choices of trade weighting (considering trade volume, source-country import and destination-country export weighting using *WTF* (NBER) data for Brazil). The difference is partly due to the fact, however, that Brazil's largest trade partners are highly flexible economies. Not weighted by trade, however, Brazil still ranks in the rigid tercile of countries.

Among the reforms, trade liberalization played a dominant role for labor-market outcomes. Multivariate regressions in Section 6 will control for sector and year effects as well as variables related to simultaneous reforms. Results will confirm the overwhelming predic-

[4]The 1988 reforms reduced the maximum working hours per week from 48 to 44, increased the minimum overtime premium from 20% to 50%, reduced the maximum number of hours in a continuous shift from 8 to 6 hours, increased maternity leave from 3 to 4 months, increased the value of paid vacations from 1 to 4/3 of the normal monthly wage, and increased the fine for an unjustified dismissal from 10% to 40% of the employer-funded severance pay account (*FGTS*). See Heckman and Pagés (2004) and Gonzaga (2003) for further details.

tive power of trade liberalization and an employer's export status for employment changes. Before an analysis or worker flows in Sections 5 and 6, however, I first turn to a conventional decomposition of employment changes in the next Section 4.

3 Linked Employer-Employee Data

Workers of particular concern for the labor-market restructuring process are prime-age male workers who typically have a low labor wage elasticity of labor supply. Most evidence of this paper nevertheless carries over to the universe of workers across gender and age groups. My restriction to prime age (of 25 to 64 years) serves to capture workers past their first entry into the active labor force.

The linked employer-employee data underlying most results reported here derive from Brazil's labor force records *RAIS* (*Relação Anual de Informações Sociais* of the Brazilian labor ministry *MTE*). *RAIS* is a nationwide, comprehensive annual census of workers formally employed in any sector (including the public sector). *RAIS* covers, by law, all formally employed workers, captures formal-sector migrants,[5] and tracks the workers over time. By design, however, workers with no current formal-sector employment are not in *RAIS*.

RAIS primarily provides information to a federal wage supplement program (*Abono Salarial*), by which every worker with formal employment during the calendar year receives the equivalent of a monthly minimum wage. *RAIS* records are then shared across government agencies and statistical offices. An employer's failure to report complete workforce information can, in principle, result in fines proportional to the workforce size, but fines are rarely 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 well above 90 percent of all formally employed workers in Brazil are covered in *RAIS* throughout the 1990s.

The full data include 71.1 million workers (with 556.3 million job spells) at 5.52 million plants in 3.75 million firms over the 16-year period 1986-2001. Every 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, and the occupation (if a worker holds multiple jobs at the same plant). Relevant worker information includes age, gender, educational attainment; job information includes tenure at the plant, occupation and the monthly average wage; plant information includes sector and municipality classifications. To facilitate tracking, *RAIS* reports formal retirements and deaths on the job. *RAIS* identifies the plant and its firm, which in turn can be linked to firm information from outside sources such as exporter data.

[5] Migration among metropolitan workers, for instance, is substantial. Among the prime-age male workers in *RAIS* with a metropolitan job in 1990, 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 elsewhere by 1995 and 27 percent by 1997. These statistics also suggest that conventional unemployment rates from household surveys may be exaggerated because migrating households are dropped from the numerator and denominator as missing, thus biasing the unemployment rate in household surveys upwards.

Table 3.1: EMPLOYMENT BY EMPLOYER'S SECTOR AND EXPORT STATUS

	Traded Goods		Nontraded Output			*Overall*[a]
	Primary	Manuf.	Comm.	Services	Other	
	(1)	(2)	(3)	(4)	(5)	(6)
			Allocation of workers, nationwide			
1990	.021	.238	.128	.280	.333	22,844
1997	.044	.195	.152	.320	.289	24,068
			Allocation of prime-age male workers, nationwide			
1990	.029	.263	.111	.284	.314	10,763
1997	.063	.221	.131	.308	.278	11,483
Nonexporter	.882	.494	.935	.937	.930	.830
Exporter	.118	.506	.065	.063	.070	.170
			Allocation of prime-age male workers, metropolitan areas			
1990	.015	.270	.104	.309	.302	5,965
1997	.024	.213	.125	.363	.275	6,060
Nonexporter	.760	.390	.887	.913	.898	.778
Exporter	.240	.610	.113	.087	.102	.222

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

Sources: Muendler (2008). *RAIS* 1990-2001, employment on Dec 31, and *SECEX* 1990-2001.

Note: Nationwide information based on 1-percent random sample, metropolitan information on 5-percent random sample. Period mean of exporter and nonexporter workforces, 1990-2001.

The samples behind results reported here chiefly derive from a list of all proper worker IDs (11-digit *PIS*) that ever appear in *RAIS* at the national level, from which a one-percent nationwide random sample and a five-percent metropolitan random sample was drawn. These randomly sampled workers are then tracked through all their formal jobs. Industry information is mostly based on the subsector *IBGE* classification (roughly comparable to the *NAICS 2007* three-digit level), which is available by plant over the full period (see Table A.1 in the Appendix for sector classifications). For the calculation of separation and reallocation statistics, a worker's separation is defined as the layoff or quit from the highest paying job.[6]

Table 3.1 shows the allocation of workers across industries in 1990 and 1997 (a detailed employment share breakdown for the *RAIS* universe can be found in Table A.1 in the Appendix). The nationwide *RAIS* records represent almost 23 million formally employed workers of any gender and age in 1990, and more than 24 million formal workers by 1997.

[6]Among the male prime-age workers nationwide, three percent of the job observations are simultaneous secondary jobs. Tables 5.1, 5.2 and 5.3 are based on the so-restricted sample, whereas all aggregate statistics, Katz-Murphy decompositions and regressions are based on the full sample. The restriction to a single job at any moment in time permits a precise definition of job separation as a layoff or quit from the highest-paying job (randomly dropping secondary jobs if there is a pay tie). Removing simultaneously held jobs does not significantly affect estimates of skill, occupation, and gender premia in Mincer (1974) regressions such as those reported in Table C.1 in the Appendix.

The bulk of Brazil's formal employment is in manufacturing, services and other industries (which include construction, utilities and the public sector), with roughly similar formal employment shares between a quarter and a third of the overall formal labor force. Commerce (wholesale and retail) employs around one in eight formal workers, and the primary sector (agriculture and mining) at most one in twenty-five formal workers.

Prime-age male workers nationwide make up slightly less than half of the total workforce in 1990 and 1997. In both years, prime-age male workers are slightly more frequently employed in the primary and manufacturing sector than the average worker of any gender and age but less frequently in commerce, services and other sectors. More than half of the *RAIS*-reported formal employment of prime-age males occurs in the six metropolitan areas of Brazil: São Paulo city, Rio de Janeiro city, Belo Horizonte, Porto Alegre, Salvador, and Recife. Compared to the nationwide average across gender and age, prime-age males in metropolitan areas are slightly less frequently employed in the primary sector, commerce, and other sectors, and somewhat more frequently employed in manufacturing and services. Overall, however, the labor allocation across sector is broadly similar across regions and gender and age groups, whereas changes over time between 1990 and 1997 are more pronounced. Between 1990 and 1997, there is a marked drop in formal manufacturing employment, which is accompanied by an increase of employment in primary sectors, commerce, and especially services. Overall, between roughly a quarter and a third of the nationwide and metropolitan prime-age male workforces are employed in traded-goods sectors, and two thirds to three quarters in nontraded-output sectors.

Table 3.2 provides a summary comparison of variables for manufacturing industries in different quintiles of comparative advantage, and between exporters and the average employer. Top comparative-advantage industries (in the highest quintile) show a higher labor turnover than the average sector with both more worker separations and more accessions, whereas exporting firms exhibit below-average turnover with both fewer worker separations and fewer accessions than average. Among the separations, reported quits play a minor role.

The average exporter is active in a sector with a slightly lower than average comparative advantage level. Similarly, there are fewer worker observations at exporters in a top comparative-advantage sector than at exporters overall. The reason is that there is a larger number of small-scale exporters in industries without comparative advantage.[7] Expectedly for a country with a history of import-substitution industrialization, Brazil's top comparative-advantage industries have lower-than-average tariffs. Comparative-advantage industries also exhibit lower import penetration. Firms in top comparative-advantage industries and exporters have larger workforces than average (85 and 326 workers more, respectively, than the average formal-sector manufacturing plant with 257 workers). The sample from *RAIS* is a random draw of workers from the formal-sector worker universe so that larger plants are over-represented. Manufacturing employment drops between 1990 and 1998, and drops faster than average in the highest-quintile advantage sectors.

To obtain labor productivity, a random extract and three-firm aggregate of the manufacturing firm survey *PIA* is used (see Appendix B). There are remarkable mean differences

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

Table 3.2: *RAIS* SUMMARY STATISTICS FOR MANUFACTURING

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

Source: Menezes-Filho and Muendler (2007). *RAIS* 1990-98 (1-percent random estimation sample), male workers nationwide, 25 to 64 years old, with manufacturing job.
Note: Statistics based on separation sample, except for accession indicator (146,787 observations in separation, 112,974 in accession sample). Sector information at subsector IBGE level. *PIA* 1986-98 for labor productivity information.

in labor productivity between an exporter and an average firm. A reason is that substantial employer heterogeneity prevails within industries, with diverse exporters and nonexporters shifting mean characteristics. Labor productivity increases between 1990 and 1998. At exporters, labor productivity is higher than average over the whole sample period, but lower than average at firms in comparative-advantage industries. Log labor productivity in 1998 exceeds log labor productivity in 1990 by 4.5 percent in the estimation sample, and by 4.7 percent at manufacturing exporters.

4 Employment Reallocation

A conventional way to measure employment reallocation is the Katz and Murphy (1992) method. The method decomposes labor demand changes into shifts between industries, associated with variations in sector sizes given sectoral occupation profiles, and within industries through changing occupation intensities. The former shifts between industries relate to the changing allocation of employment across sectors, whereas the latter shifts within industries reflect the change in relative skill intensities of occupations or alterations to the sectoral production process.

Between and within industry demand shifts. Applying the Katz and Murphy (1992) method to employment in the Brazilian formal sector over the years 1986-2001 reveals main patterns of labor-market adjustment. The decomposition into between and within sector variation indicates how two important sources of change contribute to workforce changeover. Between-industry shifts are arguably driven by changes in final-goods demands, sectoral differences in factor-nonneutral technical change, and changes in the sector-level penetration with foreign imports. Within-industry shifts can be related to factor-nonneutral technical change, factor-price changes for substitutes or complements to labor, and international trade in tasks which allocates activities along the value chain across countries.

The Katz and Murphy (1992) decomposition relates back to Freeman's (1980) manpower requirement index and is designed to measure the degree of between-industry labor demand change under fixed relative wages. The decomposition tends to understate the true between-industry demand shift in absolute terms when relative wages change. Though possibly overstating the within-industry effects, the Brazilian evidence suggests that within-industry demand changes are an important source of employment changeover in Brazil especially since 1990. Beyond the Katz and Murphy (1992) framework, I therefore offer statistics that document time variation in the occupation profile within industries, and the skill changeover within occupations.

Under the assumption that the aggregate production function is concave (so that the matrix of cross-wage elasticities of factor demands is negative semi-definite), Katz and Murphy (1992) show that an appropriate between-industry demand shift measure ΔD_k for skill group k is

$$\Delta D_k = \sum_j X_{jk} \frac{w' \, dX_j}{w' \, X_j}, \tag{4-1}$$

where X_{jk} is the employment of skill group k in industry j, w is a $k \times 1$ vector of constant wages, and dX_j and X_j are the $k \times 1$ vectors of employment changes and levels in industry j, respectively. Measure (4-1) is simply the vector of weighted sums of industry employments for each skill group k, with the weights given by the percentage changes in the overall employments in every industry j. The measure is similar to standard labor-requirement indexes Freeman (1980), only that changes are measured in efficiency units at constant wages rather than in head counts (or hours). Intuitively, skill groups that are intensely employed in expanding sectors experience a demand increase, whereas skill groups intensely employed in contracting sectors face falling demand. Under constant wages, the measure indicates whether the data are consistent with stable labor demands within sectors. Wages change, however, so that there is a bias in the measure. Katz and Murphy (1992) show that the bias is inversely related to wage changes if substitution effects dominate the employment decisions, so that measure (4-1) understates the demand increase for groups with rising relative wages.

In the Brazilian context, the formal-sector economy can be divided into 26 two-digits industries (using the subsector *IBGE* classification) and five occupations (professional & managerial occupations, technical & supervisory occupations, other white-collar occupations, skill-intensive blue-collar occupations, and other blue-collar occupations). The classification of activities into both sectors and occupations is motivated by the idea that international trade of intermediate and final goods can be understood as trade in tasks along the steps of the production chain. Using the resulting 130 industry-occupation cells, an empirically attractive version of the between-industry demand shift measure (4-1) is

$$\Delta X_k^{di} = \frac{\Delta D_k}{E_k} = \sum_i \left(\frac{E_{ik}}{E_k}\right)\left(\frac{\Delta E_i}{E_i}\right) = \frac{\sum_i \alpha_{ik}\,\Delta E_i}{E_k}, \tag{4-2}$$

where E_i is total labor input in sector-occupation cell i measured in efficiency units, and $\alpha_{ik} \equiv E_{ik}/E_i$ is skill group k's share of total employment in efficiency units in sector i in the base period. Measure (4-2) expresses the percentage change in demand for each skill group as a weighted average of the percentage changes in sectoral employments, the weights being the group-specific efficiency-unit allocations. Following Katz and Murphy (1992), I turn index (4-2) into a measure of relative demand changes by normalizing all efficiency-unit employments in each year to sum to unity. The base period is the average of the sample period from 1986 to 2001 so that α_{ik} is the share of total employment of group k in sector i over the 1986-2001 period and E_k is the average share of skill group k in total employment between 1986 and 2001.

The overall (industry-occupation) measure of demand shifts for skill group k is defined as ΔX_k^{di} from equation (4-2), where i indices the 130 industry-occupation cells. The between-industry component of this demand-shift measure is defined as the group-k index ΔX_k^{dj} from equation (4-2), where $i = j$ now indexes only 26 industries. Accordingly, the within-industry component of demand shifts is $\Delta X_k^{dw} \equiv \Delta X_k^{di} - \Delta X_k^{dj}$.

Table 4.1 presents the nationwide demand decomposition and the overall demand shifts by group of educational attainment for the economy as a whole, and separately for the traded-goods and the nontraded-output sectors. As in Katz and Murphy (1992), the percentage changes are transformed into log changes with the formula $\hat{\Delta} X_k^d = \log(1 + \Delta X_k^d)$. By construction, in the (vertical) sectoral dimension the economy-wide demand shift indices for each skill group are a weighted sum of the traded and nontraded sector indices (except for occasional rounding errors because of the log transformation), where the weights are the skill groups' shares in the sectors. In the (horizontal) time dimension, the indices are the sum of the time periods for each skill group.

The entries for overall shifts across all sectors summarize Brazil's labor-demand evolution (five first rows of column 12). Over the full period from 1986 to 2001, the least and the most skilled prime-age male workers experience a positive relative demand shift of 1 and 8 percent, respectively, whereas the three intermediate skill groups suffer a labor demand drop. This overall pattern, with demand surges at the extreme ends of the skill spectrum and drops for the middle groups, can be traced back to two overlaying developments. First, before and after the main economic liberalization episode, that is in the periods 1986-90 and 1997-2001, demand for college graduates rises by around 5 percent while demand drops for all other skill groups in 1997-2001 and for all other skill groups but high-school graduates in 1986-90. Second, during the period of economic liberalization between 1990 and 1997 the reverse labor demand change occurs, with demand for the least-educated males increasing by roughly 5 percent and dropping for college graduates by -2 percent. The demand rise for the least-educated during liberalization more than outweighs the demand drops before and after so that a net demand increase remains by 2001. For college graduates, demand surges before

Table 4.1: INDUSTRY AND OCCUPATION BASED LOG DEMAND SHIFTS, 1986-2001

(in %)	Between Industry				Within Industry				Overall Industry-Occupation			
	86-90	90-97	97-01	86-01	86-90	90-97	97-01	86-01	86-90	90-97	97-01	86-01
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Economy wide												
Illiterate or Primary Dropout	-0.1	4.7	-0.1	4.5	-2.2	-0.2	-1.0	-3.3	-2.3	4.5	-1.1	1.1
Primary School Graduate	-2.1	-0.1	-1.7	-3.9	-1.4	0.5	-1.5	-2.4	-3.6	0.4	-3.2	-6.4
Middle School Graduate	-1.9	-0.9	-0.5	-3.3	-0.1	1.5	-1.2	0.1	-2.0	0.6	-1.8	-3.1
High School Graduate	0.3	-1.7	-0.8	-2.3	1.1	0.9	0.2	2.2	1.4	-0.9	-0.6	-0.1
College Graduate	3.3	0.4	2.9	6.6	1.3	-2.4	2.5	1.4	4.6	-2.0	5.4	8.1
Traded-goods sectors												
Illiterate or Primary Dropout	-3.0	3.7	-1.7	-0.9	-0.7	-0.2	-0.2	-1.1	-3.7	3.6	-1.9	-2.0
Primary School Graduate	-3.8	-2.0	-2.4	-8.2	-0.4	0.2	-0.6	-0.7	-4.2	-1.8	-3.0	-9.0
Middle School Graduate	-3.9	-4.0	-2.6	-10.6	0.0	0.3	-0.5	-0.2	-3.9	-3.8	-3.1	-10.7
High School Graduate	-3.7	-4.4	-2.4	-10.5	0.5	-0.1	0.2	0.7	-3.2	-4.5	-2.1	-9.8
College Graduate	-3.6	-4.6	-2.1	-10.4	0.5	-0.5	1.4	1.4	-3.1	-5.1	-0.7	-8.9
Nontraded-output sectors												
Illiterate or Primary Dropout	3.6	0.4	1.9	5.9	-1.5	0.1	-0.7	-2.2	2.1	0.4	1.2	3.7
Primary School Graduate	2.8	2.7	1.5	7.0	-1.0	0.3	-0.9	-1.6	1.9	2.9	0.6	5.4
Middle School Graduate	2.3	3.3	2.3	7.9	-0.2	1.2	-0.7	0.3	2.1	4.6	1.6	8.3
High School Graduate	3.2	1.9	1.2	6.3	0.6	0.9	0.0	1.5	3.9	2.8	1.2	7.8
College Graduate	5.2	3.3	3.9	12.4	0.8	-1.8	1.4	0.4	6.0	1.5	5.3	12.8

Source: Muendler (2008). *RAIS* 1986-2001 (1-percent random sample), male workers, 25 years or older.
Note: Overall and between-industry demand shift measures for skill group k are of the form $\Delta D_k = \sum_j \alpha_{jk}(\Delta E_j / E_k)$, where α_{jk} is the average share for group k of employment in cell j over the period 1986-2001, E_j is the share of aggregate employment in cell j, and E_k is the average share of total employment of group k over the period 1986-2001 Katz and Murphy (1992). Reported numbers are of the form $\log(1 + \Delta D_k)$. In the overall measure, j indexes 130 industry-occupation cells; in the between-industry measure, $i = j$ indexes 26 industries (14 traded-goods and 12 nontraded-output sectors). The within-industry index for group k is the difference of the overall and between-industry measures. Employment is measured in efficiency units.

and after liberalization are so strong that the drop during liberalization is of little importance and a strong net demand remains by 2001. This pattern is consistent with a Heckscher-Ohlin interpretation of the specialization pattern following trade liberalization. Brazil, whose labor force is relatively low-skill abundant, experiences a shift towards low-skill intensive economic activities between 1990 and 1997—against the longer-term trend manifested before (1986-90) and after (1997-2001) by which demand for highly skilled workers increases but drops for lower-skilled workers.

Between and within decompositions, as well as a distinction of traded and nontraded sectors, lend additional support to a Heckscher-Ohlin interpretation of labor demand changes. The decomposition for all sectors (five first rows) into between-industry and within-industry changes indicates that the overall evolution is mostly driven by between-industry changes, with demand surges at the extreme ends of the skill spectrum and drops for the middle groups (column 4). In contrast, the within-industry labor demand changes favor the least skilled the least, with a demand drop of -3 percent, and the most skilled the most, with a demand increase of 1 to 2 percent for high-school educated workers and college graduates. The within-industry demand changes are almost monotonically increasing as one moves up the educational attainment ranks (column 8) in the 1986-2001 period, and would indeed monotonically increase if it were not for a within-industry drop in demand for college graduates during the liberalization period. This report will return to the within-industry demand changes with additional evidence further below. In fact, the within-industry workforce changeover will be found to reinforce a broad Heckscher-Ohlin interpretation of Brazil's experience.

A distinction by sector relates the between-industry demand evolution to differences across traded-goods industries (middle five rows) and nontraded-output industries (last five rows). In the traded-goods sectors, where trade liberalization is expected to exert its impact, Brazil experiences a salient labor demand drop—beyond -10 percent for the three more educated skill groups between 1986 and 2001. Expectedly for a low-skill abundant country, the demand drop is the strongest for the highly skilled and the weakest for the low skilled workers (column 4). Most notably, during the liberalization episode illiterate workers and primary school dropouts experience a rise in demand due to between-industry shifts, whereas more skilled workers experience demand drops of monotonically larger magnitudes as one moves up the skill ladder (column 2). The nontraded-output sectors exhibit a relatively homogeneous demand increase between 6 and 8 percent for workers with no college degree and a strong 12-percent increase for college graduates (column 4). The demand increase for the least skilled in nontraded-output sectors combined with only a slight demand drop for them in the traded-goods sectors results in an overall positive demand for the skill group from the between-industry component (column 4). Similarly, the strong demand for college graduates in nontraded-output sectors more than outweighs their demand drop in traded-goods sectors. For intermediate skill groups between these two extremes, the demand drop in the traded-goods sectors outweighs their demand increase in nontraded-output sectors and results in overall negative demand changes.

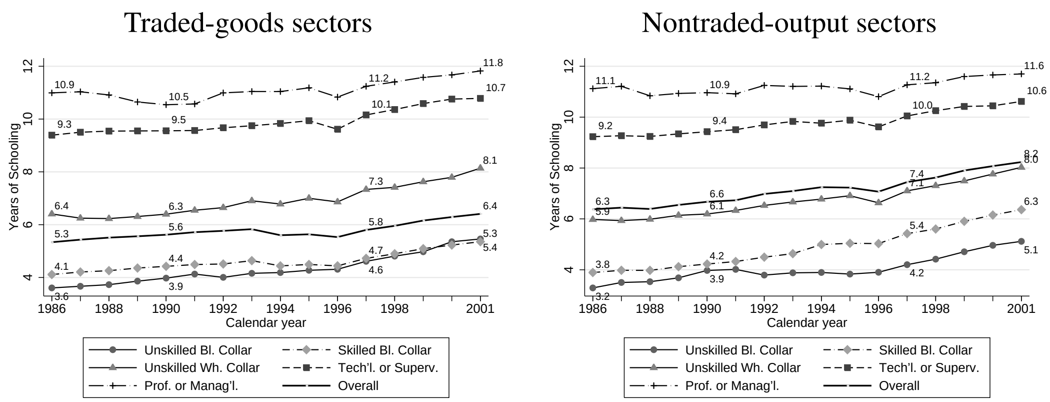
Within industries there is a clear and pronounced pattern of falling demand for the least skilled, and increasing demand for the more skilled, with monotonically stronger demand changes as one moves up the skill ranks, except only for college graduates (column 8). This pattern is similar across both traded and nontraded sectors and most time periods. The reason for the break in monotonicity at the college-graduate level (column 8) is a demand drop for this skill group during the liberalization period (column 6). A Stolper-Samuelson explanation is consistent with the outlier behavior of collage graduates during this period. Note that the Stolper-Samuelson theorem predicts wage drops for more educated workers in a low-skill abundant economy after trade reform, and Gonzaga, Menezes-Filho and Terra (2006) document that skilled earnings differentials indeed narrow over the course of the trade liberalization period. Because labor is measured in current-period efficiency units, a relative drop in wages for college educated workers tends to turn their within-industry demand index negative. With this explanation for the outlier behavior of collage graduates in view, there is a striking monotonicity in the increase in within-industry labor demand change as one moves up the skill ranks.

Within-industry employment changeovers. The demand decompositions above show a noteworthy within-industry labor demand reduction for low-skilled workers and a demand increase for high-skilled workers both in traded-goods and nontraded-output sectors. The sources of this change deserve more scrutiny. Abandoning the efficiency-unit perspective on employment in favor of counts of workers to keep wage effects separate, I turn to an assessment of labor allocation to activities by period.[8]

Figure 4.1 shows the evolution of the skill assignment by occupation over time. In both traded-goods and nontraded-output sectors, there is a marked increase across all five occupation categories in the educational attainment of the job holders. From 1986 to 2001, the mean number of years of schooling in unskilled blue-collar occupations rises from below four years to more than five years in both traded and nontraded sectors (in traded sectors schooling in unskilled blue-collar occupations even slightly exceeds the schooling in skilled blue-collar jobs by 2001). The average number of school years increases from around four to more than five years for skilled blue-collars jobs in traded sectors and to more than six years in nontraded sectors by 2001. For unskilled white-collar occupations, the average job holder's schooling goes from around six to more than eight years both in traded and nontraded goods sectors. The shift also extends to technical and supervisory positions, where the average job holder's schooling goes from less than ten to more than ten years of schooling both in traded and nontraded sectors, and to managerial positions, where mean schooling rises from eleven to almost twelve years over the period 1986-2001. These largely steady within-industry changeovers in workers' occupational assignments between 1986-2001 overlay the shorter-lived between-industry changes with much time variation across the three subperiods 1986-90, 1990-97 and 1997-01.

One might suspect that the considerable surge in schooling levels is partly due to labor supply changes such as the entry of increasingly educated cohorts of male workers into the labor force, or relatively more frequent shifts of skilled male workers from informal to formal work status over the sample period. In fact, the sector-wide average schooling level rises

[8]An efficiency-unit based analysis shows broadly the same patterns of workforce changeovers in terms of wage bills as the head-count based analysis that follows.

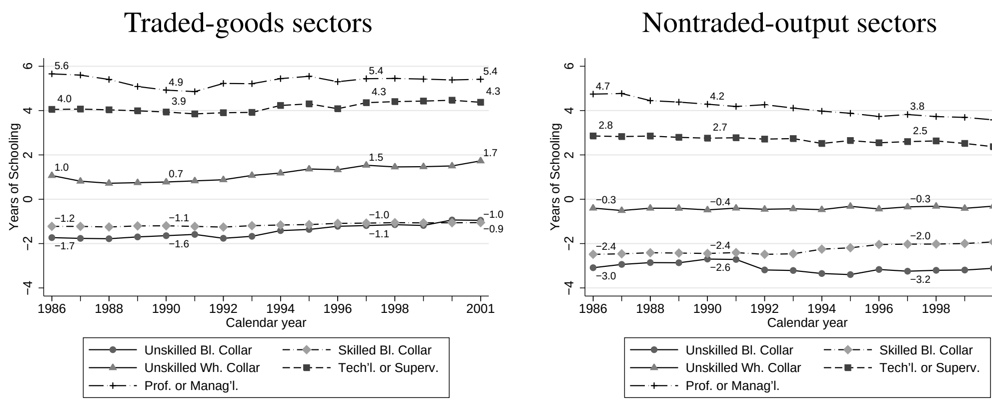


Source: Muendler (2008). *RAIS* 1986-2001 (1-percent random sample), male workers nationwide, 25 to 64 years old, with employment on December 31st.
Note: Traded-goods sectors are agriculture, mining and manufacturing (subsectors *IBGE* 1-13 and 25), nontraded-output industries are all other sectors. Mean years of schooling weighted by worker numbers within occupations.

Figure 4.1: **Schooling intensity of occupations**

from less than six to more than six years in the traded-goods sector, and in the nontraded-output sector from more than six to more than eight years (as the respective overall curves in Figure 4.1 show). To control for overall skill labor supply by sector, I extend the Katz and Murphy (1992) idea to the present context and subtract the mean annual years of schooling in a sector from the occupation-specific means in the sector. For this purpose, I consider all traded-goods industries as one sector, and all nontraded-output industries as another sector. Subtracting the annual mean years of schooling, instead of dividing by the annual total as in Table 4.1 before, preserves the cardinal skill measure of years of schooling and expresses occupation-specific skill demands as deviations from the sector-wide employment evolution in terms of years of schooling.

Figure 4.2 presents average years of schooling by occupation, less the sector-wide mean schooling across all occupations. By this measure, skill demand within every occupation category increases in the traded-goods sector since 1990: from a difference of -1.6 to -0.9 years in unskilled blue-collar occupations, from -1.2 to -1.1 years in skilled blue-collar occupations, from 0.8 to 1.7 in unskilled white-collar jobs, from 3.9 to 4.4 in technical jobs, and from 4.9 to 5.4 in professional and managerial positions. For all three white-collar occupation categories, the schooling-intensity surge beyond the sector average since 1990 is a reversal of the opposite trend prior to 1990, while schooling-intensity continually increases for blue-collar occupations in the traded sector since 1986. By construction, the persistent occupation-level increases in worker schooling since 1990 go beyond the change in the sector-wide workforce schooling. The puzzling pattern that changes beyond the sector mean are uniformly directed towards higher schooling in every single occupation since 1990 implies that there must be an employment expansion in less skill-intensive occupations—otherwise it would be impossible for every single occupation category to exhibit a faster skill-intensity increase than the average over all occupations. In contrast to the traded sector, nontraded-output industries do not exhibit the uniform pattern of schooling increases across



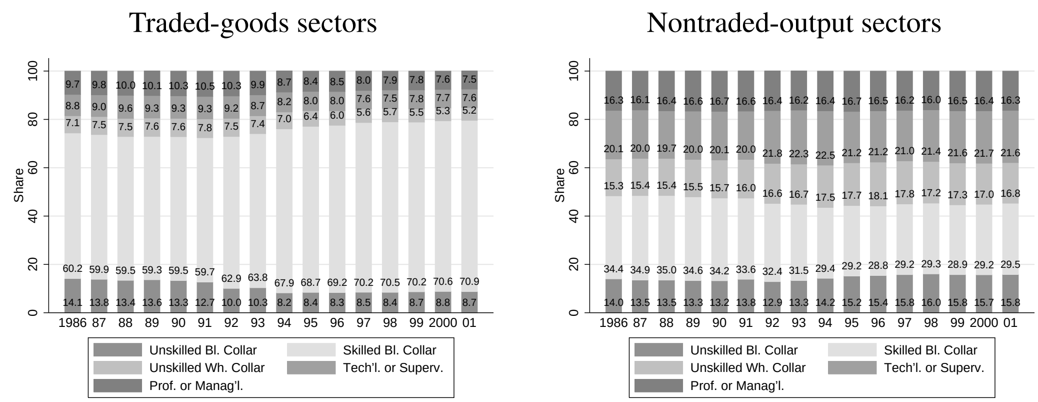
Source: Muendler (2008). *RAIS* 1986-2001 (1-percent random sample), male workers nationwide, 25 to 64 years old, with employment on December 31st.
Note: Traded-goods sectors are agriculture, mining and manufacturing (subsectors IBGE 1-13 and 25), nontraded-output industries are all other sectors. Mean years of schooling weighted by worker numbers within occupations, less mean years of schooling weighted by worker numbers across all occupations.

Figure 4.2: **Difference between schooling intensity of occupations and annual mean schooling level**

all occupations but a drop in schooling intensity in the technical and managerial occupations, and a rise in schooling intensity in skilled blue-collar occupations.

The evolution of schooling intensity in Brazil's traded-goods sector is reminiscent of a Heckscher-Ohlin interpretation as well—though not for industries but for tasks. Think of production activities in the Heckscher-Ohlin framework not as sectors but as occupations and suppose that Brazil has a relatively less schooled labor force than its main trading partners. Brazil's top five trading partners in total trade volume during the 1990s are, in descending order, the United States, Argentina, Germany, Italy and Japan. As Brazil's integration into the world economy advances, thus reinterpreted Heckscher-Ohlin trade theory predicts that Brazil increasingly specializes in less schooling intensive occupations but that Brazil employs in these expanding occupations relatively more high-skilled workers because their relative wage declines. Gonzaga et al. (2006) document that Brazil's skilled earnings differential narrows over the 1990s. Using rich linked employer-employee data that control for unobserved worker characteristics, Menezes-Filho et al. (2008) show, however, that the skill premium in wages only changes slightly between 1990 and 1997 (see Table C.1 in the Appendix). Of course, more research is required to discern this reinterpretation of classic trade theory from alternative explanations. The simultaneous schooling-intensity increase in every single occupation, above and beyond the sector mean, could also be related to factor-nonneutral technical change or factor-price changes for substitutes to labor, and not only to international trade in tasks. Yet, the prediction of reinterpreted classic trade theory that foreign trade expands less schooling-intensive occupations in Brazil's traded-goods sector is fully consistent with the data.

Figure 4.3 depicts the nationwide occupation profile within traded-goods sectors and nontraded-output sectors for the years 1986 to 2001. In traded-goods industries, skilled blue-



Source: Muendler (2008). *RAIS* 1986-2001 (1-percent random sample), male workers nationwide, 25 to 64 years old, with employment on December 31st.
Note: Traded-goods sectors are agriculture, mining and manufacturing (subsectors *IBGE* 1-13 and 25), nontraded-output industries are all other sectors. Shares based on worker numbers.

Figure 4.3: **Occupational workforce composition**

collar jobs expand markedly with the conclusion of the first wave of trade reforms between 1991 and 1993. The share of skilled blue-collar occupations increases from below 60 percent in 1990 to 68 percent in 1994 and to 71 percent by 2001. Recall from the evidence in Figure 4.1 that the average worker's schooling in both skilled and unskilled blue-collar jobs in the traded-goods sector is roughly the same. The growing importance of skilled blue-collar occupations comes at the expense of all other occupations in the traded-goods industries. At the low-skill intensity end, the share of unskilled blue-collar occupations drops from more than 13 percent in 1990 to 8 percent in 1994 (but recovers slightly to close to 9 percent by 2001). More importantly, the expansion of skilled blue-collar occupations in traded-goods sectors comes at the expense of white-collar occupations, whose total employment share drops from 27 percent in 1990 to 24 percent in 1994 and 20 percent in 2001. In the nontraded-output sectors, in contrast, it is the unskilled blue-collar occupation category that expands the fastest from 13 percent in 1990 to close to 16 percent by 2001, whereas skilled blue-collar jobs are cut back from a share of 34 in 1990 to around 29 percent by 1997. Similarly, within white-collar occupations it is again the less skill-intensive occupations that exhibit a relative gain: the share of unskilled white-collar workers rises from 16 to 18 percent between 1990 and 1995 (with a crawling scale-back to 17 percent until 2001), and the share of technical occupations increases from 20 in 1990 to 21 percent in 1995. But the share of professional and managerial positions remains roughly constant between 16 and 17 percent, thus losing in relative importance to less skill-intensive white-collar occupations.

This shift across the occupation profile towards less skill-intensive occupations permits a skill-upgrading workforce changeover, by which less skill-intensive jobs are being filled with more educated workers especially in the traded-goods sector. In practice, employers can achieve this workforce changeover in many ways. Employers can either reallocate workers across tasks in-house, or the economy can reallocate workers across firms and sectors, or there may be no reallocation for extended periods of time if employers pursue the workforce changeover by laying off less skilled workers from every occupation category in the absence of compensating rehiring within the formal sector. The latter form of workforce changeover would be associated with arguably considerable adjustment costs to the economy. As it will turn out in the next section, worker separations with little compensating rehiring elsewhere in the formal sector is prevalent.

5 Worker Reallocation Flows

Labor-demand decompositions so far have shown that there are two main components to the observed workforce changeover in Brazil over the sample period. First, there is a labor demand shift towards the least and the most skilled male workers, which can be traced back to relatively weaker declines of traded-goods industries that intensely use low-skilled labor and to relatively stronger expansions of nontraded-output industries that intensely use higher-skilled labor. Second, there is a within-industry shift towards longer-schooled workers, associated with a skill-upgrading of all occupations in traded-goods industries.

The conventional decomposition leaves unaddressed, however, how the workforce changeover come about. To analyze how employers achieve the observed workforce changeover, actual worker flows need to be observed and comprehensive linked employer-employee data are required. Linked employer-employee data for Brazil's economy trace individual workers across their jobs within plants, across plants within sectors, and across firm types and sectors in Brazil's formal sector.

Reallocations across tasks. Employers may choose to reallocate workers across tasks in-house. For this purpose, define an in-house job change as a change in employment between an occupation at the *CBO* base-group level to another base-group occupation. The 354 *CBO* base groups roughly correspond to the 4-digit *ISCO-88* occupations at the unit-group level.[9] Table 5.1 shows both continuing and displaced workers and tracks the workers through jobs at the annual horizon between 1986 and 1997. The task assignment pattern is remarkably stable both before and after trade liberalization. Between 86 and 87 percent of formal-sector prime-age male workers remain in their job at the same employer. Only between 1 and 2 percent of the workers are assigned to new occupations within the same plant. Less than one percent of the workers switches plants within the same firm. Between 7 and 9 percent of the workers change employing firm at the annual horizon. So, the bulk of successful reallocations does not take place on internal labor markets but across firms. Reallocations between exporters and nonexporters and across sectors will be reported below. The remaining 3 to 4 percent of workers (not reported in Table 5.1) are unaccounted. Those failed reallocations will also be revisited shortly. Overall, the stable and minor percentages of occupation and plant reassignments within employers suggest that the observed workforce changeovers, documented in the preceding section, are not achieved through job reassignments in internal labor markets.

[9] For a description of the Brazilian occupation classification system *CBO* and a mapping to *ISCO-88*, see Muendler, Poole, Ramey and Wajnberg (2004).

Table 5.1: ANNUAL OCCUPATION CONTINUATIONS AND TRANSITIONS 1986-97

Year t	1986	1988	1990	1992	1994	1996
Year $t+1$	(1)	(2)	(3)	(4)	(5)	(6)
Employed						
in same occupation	.867	.859	.864	.859	.850	.856
at same establishment in new occupation	.018	.018	.019	.020	.020	.013
at same firm but new establishment	.007	.006	.006	.007	.006	.005
at new firm	.079	.084	.074	.078	.087	.083

Source: Muendler (2008). *RAIS* 1986-97 (1-percent random sample), male workers, 25 years or older.
Note: Frequencies based on last employment of year (highest paying job if many); continuations at same firm exclude continuations at same plant. Occupations are defined at the CBO 3-digit base-group level with 354 categories, which roughly correspond to the 4-digit ISCO-88 unit-group level.

Reallocations across firms and sectors. Between 1990 and 1998, around 6 percent of the formal-sector workforce nationwide is employed at primary-sector nonexporters, one percent at primary-sector exporters, 11 percent at manufacturing nonexporters, and 12 percent at manufacturing exporters. The remaining seventy percent of the workforce are employed in the nontraded sector. Looking beyond internal labor markets, linked employer-employee data permit an investigation into whether and how the relative expansion of certain traded-goods industries, in the wake of an overall decline of the traded-goods sector, is associated with reallocations of individual workers across firms and sectors. To capture differences in the labor demand responses across subsectors and firms within the traded-goods sector, the following tabulations track individual workers across exporting and nonexporting employers in the primary and manufacturing industries.

Table 5.2 shows worker transitions between firms and sectors over the first year after trade reform, between their last observed formal-sector employment in 1990 and their last observed formal-sector employment in 1991. Only workers who experience a separation from their last employment of the year are included in the transition statistics. Trade theory might lead one to expect a shift of displaced workers from nonexporting firms to exporters following trade reform. Although manufacturing exporters are only about five percent of firms during the 1990s, they employ about half the manufacturing workforce. The dominant share of successful reallocations of former nonexporter workers within the traded-goods industries, however, is to nonexporters again. Among the former nonexporter workers displaced from primary-sector employment, close to 11 percent are rehired at primary nonexporters and 10 percent at manufacturing nonexporters, but less than two percent shift to exporters. Among the former nonexporter workers in manufacturing, 19 percent move to manufacturing nonexporters and 7 percent no manufacturing exporters, and a very small share to primary-sector firms. Former exporter workers, in contrast, mostly transition to new formal-sector jobs within the sector of displacement and are roughly equally likely to find reemployment at an exporter or a nonexporter. These patterns suggest that reallocations within the traded goods sectors are mostly intra-sector reallocations from exporter to exporter and from nonexporter to nonexporter—contrary to what classic trade theory with full employment and only traded

Table 5.2: YEAR-OVER-YEAR FIRM AND SECTOR TRANSITIONS, 1990-91

To:	Primary		Manufacturing		Nontraded	*Failure*	*Total*
	Nonexp.	Exp.	Nonexp.	Exp.			
From: (in %)	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Primary Nonexporter	10.7	.7	10.3	1.2	40.3	36.8	100.0
Primary Exporter	6.7	6.7	3.3	3.3	45.0	35.0	100.0
Manufact. Nonexporter	1.4	.1	19.3	7.2	34.9	37.1	100.0
Manufact. Exporter	1.2	.1	14.5	15.5	33.5	35.2	100.0
Nontraded	1.3	.0	5.4	2.4	54.8	36.0	100.0
Failure	2.9	.3	13.2	5.6	78.0	.	100.0
Total	2.1	.2	10.1	4.8	59.7	23.2	100.0

Source: Muendler (2008). *RAIS* 1990-91 (1-percent random sample), male workers nationwide, 25 to 64 years old. *SECEX* 1990-91 for exporting status.
Note: 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 in earlier job.

goods might lead us to expect.

In the initial year after trade reform, between one third and two-fifths of displaced traded-sector workers with a successful reallocation end up in nontraded-sector jobs. An equally large fraction, however, fails to experience a successful reallocation to any formal-sector job within the following calendar year (retirements, deaths, and workers at or past retirement age are excluded from the displaced worker sample).[10] Of the workers with a failed reallocation before year-end 1990, by far the largest fraction (of 78 percent) with a successful reallocation by year-end 1991 finds employment in the nontraded-sector. In summary, at the time of the largest impact of trade liberalization in 1990-91, traded-goods industries exhibit little absorptive capacity for displaced workers compared to nontraded-output industries and compared to the prevalence of failed transitions out of the formal sector. Among those failed reallocations can be transitions to informal work, unemployment, or withdrawals from the active labor force, which are not directly observed in the *RAIS* records.[11]

In comparison, Table 5.3 tracks annual transitions six years after the beginning of trade liberalization and three years after its conclusion. By 1996-97, more firm and sector reallocations from the primary sector are directed to jobs within the traded-goods sector. In the manufacturing sector, however, the dominant destination sector of displaced workers remains the nontraded sector in 1996-97, both for workers from exporters and for workers from nonexporters. As in the initial period 1990-91, in 1996-97 former nonexporter workers most frequently find reemployment at nonexporter firms, and former exporter workers are

[10]The slightly smaller unaccounted percentage in Table 5.1 compared to the reallocation failure rates in Tables 5.2 and 5.3 is largely due a restriction of the initial sample to workers with comprehensive occupation information in Table 5.1.

[11]For evidence on those work status transitions using household survey data, see Menezes-Filho and Muendler (2007).

Table 5.3: YEAR-OVER-YEAR FIRM AND SECTOR TRANSITIONS, 1996-97

To:	Primary		Manufacturing		Nontraded	*Failure*	*Total*
	Nonexp.	Exp.	Nonexp.	Exp.			
From: (in %)	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Primary Nonexporter	32.1	2.5	6.0	2.9	15.4	41.1	100.0
Primary Exporter	17.1	13.0	6.5	3.3	18.7	41.5	100.0
Manufact. Nonexporter	5.6	.4	18.9	6.5	32.1	36.5	100.0
Manufact. Exporter	7.2	.7	12.1	13.9	27.3	38.8	100.0
Nontraded	1.3	.2	3.8	2.0	55.8	36.9	100.0
Failure	8.9	.7	12.2	6.1	72.1	.	100.0
Total	6.5	.6	8.8	4.7	56.9	22.5	100.0

Source: Muendler (2008). *RAIS* 1996-97 (1-percent random sample), male workers nationwide, 25 to 64 years old. *SECEX* 1996-97 for exporting status.
Note: 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 in earlier job.

roughly equally likely to find reemployment at exporter and nonexporter firms in manufacturing but less likely to transition to an exporter in the primary sector. By 1996-97, an even larger fraction of displaced primary-sector workers than in 1990-91 fails to experience a successful formal-sector reallocation and a roughly equally large share of former manufacturing workers as in 1990-91 fails to find a formal-sector job within the following calendar year.

Together with the evidence on infrequent task reassignments in-house, these labor-market transitions suggest that the observed workforce changeovers from the preceding section are neither achieved through worker reallocations within employers nor are they brought about by labor reallocations across employers and sectors. By exclusion, the remaining explanation is that formal-sector employers in the traded-goods industries shrink their workforces by dismissing less-schooled workers more frequently than more schooled workers while the thus displaced workers fail to find reemployment at least at the annual horizon. In the aggregate, the lacking traded-sector reallocations result in a considerable decline of formal manufacturing employment from 26 to 22 percent (Table 3.1). The simultaneous expansion of nontraded-output industries can partly be driven by a long-term shift from primary to manufacturing to services activities in the economy, or by trade liberalization if fast productivity change reduces manufacturing employment in favor of non-traded sector employment, or by Brazil's overvalued real exchange rate during the sample period, or by foreign direct investment (FDI) flows in the wake of Brazil's concomitant capital-account liberalization and privatization programme, or by a combination of these changes. The next section turns to the predictive power of these competing explanations and their associated variables, using linked employer-employee data at the job level.

6 Trade-Induced Worker Separations and Accessions

Employers adjust workforces through worker separations and accessions. A separation is defined as a worker's quit or layoff from the last formal employment in the calendar year. Among the separations, quits are infrequent compared to layoffs (Table 3.2).[12] Conversely, an accession is defined as a worker's hiring into the first formal employment in the calendar year. Separations in turn burden, and accessions unburden, the pool of workers to be reallocated.

To understand determinants of labor reallocation in the formal sector, regression analysis can simultaneously condition on industry, plant, job and worker characteristics as explanatory variables for separations and accessions. Consider the probability that an employer-employee match is terminated (a separation) or is formed (an accession), conditional on a worker-fixed component α_i that is observable to the employer and the worker:

$$Pr\left(\sigma_{i,t} \middle| \mathbf{x}_{i,t}, \mathbf{y}_{J(i),t}, \mathbf{z}_{S(J(i)),t}\right) = \frac{\exp\{\mathbf{z}_{S(J(i)),t}\beta_z + \mathbf{y}_{J(i),t}\beta_y + \mathbf{x}_{i,t}\beta_x + \alpha_i + \alpha_t\}}{1 + \exp\{\mathbf{z}_{S(J(i)),t}\beta_z + \mathbf{y}_{J(i),t}\beta_y + \mathbf{x}_{i,t}\beta_x + \alpha_i + \alpha_t\}}, \quad \text{(6-1)}$$

where $\sigma_{i,t}$ denotes the binary outcome (accession or not, separation or not) for worker i at time t. $\mathbf{z}_{S(J(i)),t}$ is a vector of sector-level covariates of the worker's displacing or hiring sector $S(J(i))$, including a sector-fixed effect in some specifications; $\mathbf{y}_{J(i),t}$ is a vector of plant-level covariates of worker i's displacing or hiring plant $J(i)$; $\mathbf{x}_{it}$ is a vector of covariates that are worker, job or match specific; β_z, β_y, β_x are coefficient vectors; α_i is the worker-fixed effect and α_t a year effect. There is an unobserved error to terminations and formations of employer-employee matches. For theoretical consistency with random shocks to employer-employee matches, the disturbance is assumed to be logistic and independent across matches. This conditional logit model (6-1) is fit 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. Consequently, education categories are dropped from the worker characteristics vector but educational workforce composition shares are kept among the plant-level regressors. When inferring separations and accessions in this and subsequent sections, transfers across plants within the same firm, as well as retirements and reported deaths on the job are excluded.

Table 6.1 presents conditional logit estimates of separations from formal manufacturing jobs, where the conditioning removes worker-fixed effects (worker-FE logit) and year effects. For comparison, the first five columns present regressions without sector-fixed effects so that sector-specific variables such as comparative advantage (which varies little over time) can be kept among the regressors. Separations are significantly more frequent in sectors with a stronger comparative advantage and at exporters—contrary to predictions of standard trade theory. Elevated product tariffs predict lower separation rates from formal jobs (though only

[12] Separations are treated as a single category for regression analysis, where no marked differences between quits and layoffs for trade-related predictors can be detected.

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 (Corden 1966, Anderson 1998). Similarly, additional import penetration predicts significantly higher displacement odds. When including observed market penetration with imports to proxy for changing non-tariff barriers and all earlier trade related predictors, point estimates and statistical significance of coefficients are hardly affected as the specification is gradually enriched (moving from column 1 to column 6). FDI inflows into the sector predict a statistically significant reduction in displacement rates. The sectoral real exchange and the Herfindahl concentration index have no significant predictive power after conditioning on year effects.

When year indicators are excluded from the regression (column 5), comparative advantage and exporting status become even stronger predictors of displacements. Tariffs and import penetration coefficients now also reflect the effect of reducing trade barriers over time and predict that reduced barriers both at the input and the output margin, and the arrival of additional imports, are associated with more worker separations. Using further controls—such as the inflation rate in addition to sectoral price levels behind the real exchange rate, FDI stocks in addition to FDI flows, and controls for privatization and outsourcing—beyond the large set of sector- and firm-level variables that already control for time-varying changes to the competitive environment does not change coefficients in important ways.

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

Before discussing plant and worker-level variables, turn to the opposite margin: Table 6.2 presents conditional logit estimates of accessions into formal manufacturing jobs, controlling for worker-fixed accession effects. Mirroring the signs from separation regressions, accession rates are lower in sectors with stronger comparative advantage, when other trade-related variables are controlled for (column 4). The coefficient is not statistically significant at conventional levels in this regression (but will become statistically significant when controlling for higher-order interactions between trade variables in Table 7.3). Exporters exhibit significantly lower accession rates, mirroring their higher separation rates. Elevated product tariffs predict significantly more accessions, mirroring the sign from separation regression, whereas higher intermediate-input tariffs predict significantly fewer accessions, also mirroring the sign from separation regression. Import penetration has no statistically significant effect, and neither does the real exchange rate. FDI inflows are associated with significantly more accessions and more concentrated manufacturing industries exhibit fewer accessions.

When year effects are omitted (column 5), comparative advantage and exporting status become even stronger predictors of reduced accessions. Tariffs and import penetration coef-

Table 6.1: WORKER-FIXED EFFECT LOGIT ESTIMATION OF SEPARATIONS

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

Source: Menezes-Filho and Muendler (2007). *RAIS* 1990-98 (1-percent random sample), male workers nationwide, 25 to 64 years old, with manufacturing job.
Note: Separations exclude transfers, deaths, and retirements. Reference observations are employments with no reported separation in a given year. Sector information at subsector IBGE level. Professional or managerial occupations and skilled blue collar occupations (not reported) not statistically significant at five-percent level. Robust standard errors in parentheses: * significance at ten, ** five, *** one percent.

Table 6.2: WORKER-FIXED EFFECT LOGIT ESTIMATION OF ACCESSIONS

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

Source: Menezes-Filho and Muendler (2007). *RAIS* 1990-98 (1-percent random sample), male workers nationwide, 25 to 64 years old, with manufacturing job.
Note: Accessions exclude transfers. Reference observations are employments with no reported accession. Sector information at subsector IBGE level. Robust standard errors in parentheses: * significance at ten, ** five, *** one percent.

ficients now also reflect the effect of reducing trade barriers over time. Lower input tariffs, which tend to make competition less fierce, predict more accessions. Lower output tariffs and the arrival of additional imports, which tend to make competition more fierce, are associated with fewer accessions. When conditioning on both year and sector effects (column 6), the largely time-invariant a sector-specific comparative advantage variable does expectedly not turn significant, whereas coefficients for all other trade regressors increase in absolute value (compared to column 4) and remain or become highly significant. As for separations, this report therefore bases much of the subsequent discussion on the more conservative estimates without sector effects.

Larger manufacturing plants offer more employment stability: they displace significantly fewer (Table 6.1) and they hire significantly fewer workers (Table 6.2). Plants with less educated workforces and more blue-collar jobs separate from workers significantly less frequently and hire significantly more frequently. Interestingly, workers with a longer tenure at the plant and longer labor-market experience suffer significantly more frequent separations at the separation margin. This fact is consistent with the hypothesis that Brazilian firing costs, which proportionally increase with tenure, lead employers to shorten tenure through displacement. Workers in occupations of intermediate skill intensity experience significantly fewer separations, and workers are significantly less likely to be hired into high-skill intensive manufacturing occupations (with a monotonic drop in accession odds as an occupation's skill intensity increases). Year effects are significant at the one-percent level and show both a strictly monotonic increase in manufacturing separations and a strictly monotonic drop in manufacturing accessions.

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

The evidence so far shows that Brazil's trade reform predicts salient changes to worker separations and accessions. But neither comparative-advantage sectors nor exporters exhibit the expected labor absorption; they separate from their workers significantly more frequently than other sectors and firms. Exporters also hire significantly less frequently. There are empirical concerns for these predictions of worker flows:r the potential simultaneity of trade policies and exporting status, and the relevance of Brazil's concomitant reforms.

Trade exposure and exporting status. Despite the apparently exogenous nature of trade reform for individual employers—the enactment by decree on president Collor's first day in office surprises politicians and businesses alike—, the reduction in tariff dispersion gives rise to a simultaneity concern. By design, initially highly protected sectors face the largest product tariff declines. Similarly, market penetration with foreign inputs possibly responds to Brazilian labor-market conditions. In a two-stage least squares approach, it is therefore important to 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. So export status should also be predicted with instrumental variables in a two-stage least squares approach.

To construct instruments for export demand, one can consider seven broad destination regions of Brazil's exports, calculate the destination-region imports from other source countries than Brazil, and weight the destinations' importance for Brazil's industries using Brazilian exports in 1990. These foreign demand proxies vary by sector and year. In addition, the nominal U.S. dollar exchange rate and sector-level producer-price indices in the United States and the EU are valid instruments. Nominal exchange rate movements are largely unpredictable, and current foreign producer price levels in industrialized economies are arguably unrelated to the termination and formation of job matches in Brazil. To check for potential sign reversals and assess the magnitude of possible simultaneity bias, it is useful to resort to linear fixed-effects regressions of separation and accession indicators on the same predictors as in the preceding section:[13]

$$\sigma_{i,t} = \hat{\mathbf{z}}_{S(J(i)),t}\beta_z + \hat{\mathbf{y}}_{J(i),t}\beta_y + \mathbf{x}_{i,t}\beta_x + \alpha_i + \alpha_t + \epsilon_{i,t}, \tag{6-2}$$

where $\sigma_{i,t} \in \{0,1\}$ denotes the binary outcome (accession or not, separation or not) for worker i at time t, and regressor and coefficient vectors are as in (6-1). There is an unobserved error $\epsilon_{i,t}$ to the termination and formation of employer-employee matches. It is assumed to be normally distributed and independent across employer-employee matches. For the two-stage least squares approach, 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 (6-2). Turning to linear regression has the additional benefit that the estimation sample includes workers with no change in employment; their worker-fixed effect is separately identified through time variation of other predictors at the same employer. The change in estimation sample affords an additional robustness check.

On the first stage, export status, product and input tariffs, and import penetration are regressed on the instrumental variables, weighting the regressions by employment observations in the separation and accession samples. Table 6.3 shows the results by sample, except for the input-tariff estimates, which are similar to product-tariff estimates. There is no evidence of weak instruments: F statistics from joint significance tests on the instruments vary between 13 and 14,000. Almost invariably, the instruments are statistically significant predictors at the one-percent level.[14] Expectedly, higher producer prices in the United States and Europe, as well as a weaker Brazilian currency, predict significantly more frequent exporting status. Employment-weighted exporting status is more frequent in sectors with weaker comparative advantage, as documented in the data Section 3 before, because there is a larger number of small-volume exporters in the low-advantage sectors.

[13]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).

[14]Experimenting with labor productivity in the initial year 1990 as a candidate firm-level instrument in the subsample of *PIA* firms and performing over-identification tests shows that the validity of labor productivity is rejected.

Table 6.3: FIRST-STAGE PREDICTIONS WITH SECTOR EFFECTS

	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	-2.705	-2.669	.977	-1.922	-2.864	.887
	(1.252)**	(.121)***	(.055)***	(1.613)	(.143)***	(.066)***
World imports CEE	15.389	-79.896	-7.277	5.738	-73.858	-7.262
	(4.919)***	(.477)***	(.217)***	(6.379)	(.566)***	(.262)***
World imports LAC	-10.046	14.864	3.370	-7.802	14.012	3.352
	(1.454)***	(.141)***	(.064)***	(1.939)***	(.172)***	(.080)***
World imports NAM	3.973	-5.427	-.625	4.098	-4.023	-.803
	(.874)***	(.085)***	(.039)***	(1.144)***	(.101)***	(.047)***
World imports ODV	2.394	-3.272	4.574	3.739	-3.012	4.968
	(.961)**	(.093)***	(.042)***	(1.288)***	(.114)***	(.053)***
World imports OIN	17.934	9.786	-1.810	16.801	7.407	-1.008
	(1.921)***	(.186)***	(.085)***	(2.431)***	(.216)***	(.100)***
World imports WEU	9.690	2.455	-.741	9.914	1.955	-.757
	(.659)***	(.064)***	(.029)***	(.834)***	(.074)***	(.034)***
USD Exch. Rate	.078	-.158	.035	.091	-.178	.027
	(.027)***	(.003)***	(.001)***	(.035)***	(.003)***	(.001)***
PPI Idx. EU	.463	-.977	.149	.734	-.964	.111
	(.119)***	(.011)***	(.005)***	(.149)***	(.013)***	(.006)***
PPI Idx. NAM	-.058	.482	-.364	.059	.487	-.440
	(.114)	(.011)***	(.005)***	(.149)	(.013)***	(.006)***
Exogenous covariates						
Balassa Comp. Adv.	.019	-.015	.008	.011	-.014	.007
	(.006)***	(.0006)***	(.0003)***	(.008)	(.0007)***	(.0003)***
FDI Flow (USD billion)	-.004	.008	-.001	-.005	.007	-.0001
	(.003)	(.0003)***	(.0001)***	(.004)	(.0003)***	(.0002)
Herfindahl Index (sales)	.263	.037	-.468	.208	.002	-.448
	(.084)***	(.008)***	(.004)***	(.107)*	(.010)	(.004)***
Log Employment	.052	.0003	-.0004	.050	.0006	-.0005
	(.002)***	(.0002)*	(.00007)***	(.002)***	(.0002)***	(.00007)***
Share: Middle School or less	-.169	.001	-.003	-.183	.0005	-.002
	(.016)***	(.002)	(.0007)***	(.017)***	(.002)	(.0007)***
Share: Some High School	-.059	.0007	-.0004	-.092	-.002	.001
	(.019)***	(.002)	(.0008)	(.021)***	(.002)	(.0009)
Share: White-collar occ.	.073	.001	-.003	.068	.0007	-.003
	(.010)***	(.001)	(.0004)***	(.012)***	(.001)	(.0005)***
F statistic (IV)	9.832	7,885.4	6,918.1	9.851	6,545.8	6,013.2

Sources: Menezes-Filho and Muendler (2007). *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.
Note: Weighted regressions using worker-sample observations (as in Table 6.1 for separations, Table 6.2 for accessions), controlling for year and sector effects. Annual sector-weighted world imports, coefficients rescaled to imports in USD trillion. Robust standard errors in parentheses: * significance at ten, ** five, *** one percent.

Table 6.4: 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.	.010	-.006	-.012	-.067	-.007	-.013
	(.045)	(.005)	(.009)	(.048)	(.004)	(.007)**
Exporter Status	.293	.037	.053	-.438	-.049	.123
	(.028)***	(.003)***	(.178)	(.027)***	(.003)***	(.135)
Product Market Tariff	-2.197	-.295	-.567	1.822	.162	.255
	(.469)***	(.042)***	(.077)***	(.498)***	(.038)***	(.070)***
Intm. Input Tariff	.265	.463	.754	-2.954	-.216	-.201
	(.610)	(.065)***	(.109)***	(.750)***	(.059)***	(.101)**
Import Penetration	9.014	-.068	-.266	1.764	.362	.454
	(.463)***	(.054)	(.103)***	(.665)***	(.049)***	(.089)***
Obs.	145,408	293,353	293,353	112,974	293,124	293,124

Source: Menezes-Filho and Muendler (2007). *RAIS* 1990-98 (1-percet random sample), male workers nationwide, 25 to 64 years old, with manufacturing job.
Note: Separations and accessions exclude transfers, deaths, and retirements. Reference observations are employments with no reported separation or accession in a given year. Sector information at subsector IBGE level. Estimates in column 1 and 4 repeat column 4 in Tables 6.1 and 6.2. Further regressors (not reported): Year indicators, sector indicators, sector, plant and worker covariates. Robust standard errors in parentheses: * significance at ten, ** five, *** one percent.

Turn to the second stage. For ease of comparability, Table 6.4 restates conditional logit estimates for separations and accessions in columns 1 and 4 (from Tables 6.1 and 6.2, column 4). Table 6.4 contrasts those earlier estimates with linear worker-fixed effects regressions without (columns 2 and 5) and with instrumental-variable predictions (columns 3 and 6). The estimation samples for the linear worker-fixed effects models are substantially larger because workers with no transition remain in the sample. When instrumenting, there is not a single sign reversal in the potentially simultaneity-afflicted coefficients—export status, tariffs, and import penetration (comparing columns 2 and 3, and columns 5 and 6). Instrumentation overwhelmingly reinforces at the one-percent significance level that comparative-advantage sectors and exporters exhibit more separations, and exporters exhibit fewer accessions. Several coefficients on tariffs and import penetration lose significance at common levels under instrumental-variable fixed-effects regressions (columns 3 and 6) but never exhibit a sign reversal.[15] So instrumentation in a linear probability model corroborates the main hypothesis regarding 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.

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

Economic change and reforms. The Brazilian economy undergoes a series of concomitant economic transformations during the sample period, including technological changes, the intensified outsourcing of service jobs, surging foreign direct investment inflows and policy shifts such as macroeconomic stabilization, capital-account liberalization, and privatization.[16] The accession and separation regressions so far control for sector and year covariates including sectoral real exchange rates, Herfindahl sales concentration indices, foreign direct investment inflows, and sector and year fixed effects. It remains to turn to economic changes and policies that may affect estimates at the level of the plant, job, worker or employer-employee match in specification (6-1).

If skill-biased technological change systematically interacts with the effect of trade reform on labor turnover, trade reform expectedly covaries with labor turnover differently for workers with different skills. To check for this heterogeneity, one can run specification (6-1) separately for young workers with less than ten years of potential labor-market experience, and for workers with primary schooling and some college education. Table 6.5 redisplays in column 1 the conditional logit estimates for separations and accessions on the full sample. Estimates for the skill subsamples follow in columns 2 through 4. Coefficient estimates for separations and accessions are strikingly similar across the samples. No sign changes. Import penetration predicts a stronger effect on young workers' separations and accessions and, surprisingly, implies that trade integration predicts more frequent separations and significantly less frequent accessions for young workers. This also suggests that, if anything, the restriction of the regression sample to prime-age workers biases trade effects against the main hypothesis. Magnitudes of the tariff and import-penetration coefficients significantly increase for more educated workers, but are statistically indistinguishable for comparative advantage and export status. Statistical significance is lost in some cases in the smaller college-educated worker subsample. There is, surprisingly, no strong evidence that skill-biased labor-demand changes systematically interact with the effect of trade reform on separations and accessions.

The privatization of state-owned businesses and the progressing outsourcing of service jobs to specialized suppliers can affect separations and accessions. If privatization and outsourcing covary with the trade regime and labor turnover in systematic ways, they potentially lead to erroneous attribution. The ownership status of a plant is observable in *RAIS* since 1995, when the federal government started to pursue privatization on a larger scale. To control for privatization, 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 (6-1). As column 5 in Table 6.5 shows, coefficient estimates on the trade-related variables exhibit no statistically significant change, and the ownership-status itself is not a statistically significant predictor. Define a job as susceptible 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 (6-1) results in no statistically significant coefficient change (column 6). Jobs susceptible to outsourcing exhibit a statistically significant reduction in accession odds. There is, in summary, no evidence that

[16] Labor-market institutions were altered preceding trade reform and their industry-specific impact is controlled for with sector-fixed effects (Tables 6.1 and 6.2, column 6).

Table 6.5: ALTERNATIVE WORKER-EFFECT LOGIT SPECIFICATIONS

	Cdl. logit	Young worker	Primary school	College educ.	Privat. control	Outsrc. job ind.
	(1)	(2)	(3)	(4)	(5)	(6)
			SEPARATIONS			
Balassa Comp. Adv.	.010	.856	-.135	.099	-.138	-.091
	(.045)	(.543)	(.055)**	(.338)	(.050)***	(.050)*
Exporter Status	.293	.361	.299	.284	.283	.284
	(.028)***	(.239)	(.033)***	(.145)**	(.028)***	(.029)***
Product Market Tariff	-2.197	-4.334	-2.200	-2.282	-2.809	-2.320
	(.469)***	(4.556)	(.556)***	(2.378)	(.492)***	(.479)***
Intm. Input Tariff	.265	4.686	5.126	4.216	5.696	5.188
	(.610)	(7.686)	(.859)***	(4.269)	(.768)***	(.756)***
Import Penetration	9.014	2.833	3.655	.425	3.057	3.320
	(.463)***	(5.490)	(.770)***	(3.089)	(.638)***	(.643)***
addl. regressor(s)					-7.924	-.007
					(1.718)***	(.037)
Obs.	145,408	2,897	110,831	7,498	145,408	143,536
Pseudo R^2	.144	.395	.162	.249	.152	.153
			ACCESSIONS			
Balassa Comp. Adv.	-.067	-.284	-.041	-.726	-.037	-.087
	(.048)	(.517)	(.054)	(.343)**	(.049)	(.049)*
Exporter Status	-.438	-.536	-.421	-.775	-.438	-.437
	(.027)***	(.219)**	(.031)***	(.143)***	(.027)***	(.027)***
Product Market Tariff	1.822	-1.785	1.865	.938	2.044	1.638
	(.498)***	(3.803)	(.576)***	(2.426)	(.502)***	(.503)***
Intm. Input Tariff	-2.954	-3.380	-2.849	-1.833	-3.204	-2.773
	(.750)***	(6.677)	(.854)***	(3.975)	(.753)***	(.758)***
Import Penetration	1.764	-.588	1.059	3.557	1.848	1.621
	(.665)***	(5.041)	(.799)	(3.333)	(.668)***	(.670)**
addl. regressor(s)					6.150	-.095
					(1.840)***	(.033)***
Obs.	112,974	2,752	86,468	4,786	112,974	110,985
Pseudo R^2	.042	.238	.043	.094	.042	.041

Source: Menezes-Filho and Muendler (2007). *RAIS* 1990-98 (1-percent random sample), male workers nationwide, 25 to 64 years old, with manufacturing job.
Note: Separations exclude transfers, deaths, and retirements. Reference observations are employments with no reported separation in a given year. Sector information at subsector IBGE level. Estimates in column 1 repeat column 4 in Tables 6.1 and 6.2. Further regressors (not reported): Year indicators, sector indicators, sector, plant and worker covariates. Robust standard errors in parentheses: * significance at ten, ** five, *** one percent.

simultaneous economic changes and concomitant reforms systematically alter the effect of trade reform on separations and accessions.

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

7 Explanations of Reallocation Flows

A strong candidate explanation for the reverse labor flows away from comparative-advantage sectors and away from exporters is the endogenous change in productivity. Several case studies have documented for various countries in the context of trade liberalization episodes and other structural reforms that within-firm productivity rises in response to the removal of trade protection (e.g. Levinsohn 1993, Hay 2001, Pavcnik 2003, Schor 2003, Eslava, Haltiwanger, Kugler and Kugler 2004, Fernandes 2007, Muendler 2004). If trade triggers faster productivity growth at exporters and in comparative-advantage industries because for these firms and industries larger market potential offers stronger incentives to improve efficiency, and if productivity increases faster than production, then labor flows away from comparative-advantage sectors and away from exporters. Production, and market shares, increase less than proportionally with productivity if the elasticity of demand is less than unity in absolute value. As a result, output shifts to more productive firms but labor does not.

Firm-level labor productivity. Exporters are more productive than nonexporters, as Table 3.2 has documented. To compare the relative importance of a firm's exporter status and labor productivity for separations and accessions, it is desirable to include a measure of firm-level labor productivity in specification (6-1). For this purpose, I use the subsample of *RAIS* firms that are also covered in the *PIA* manufacturing survey, for which firm-level labor productivity is inferrable. This link reduces the number of observations markedly. Moreover, confidentiality requirements only allow me to use randomly combined three-firm cells from *PIA*, resulting in a loss of efficiency.

Table 6.4 redisplays conditional logit estimates for separations and accessions on the full sample in columns 1 and 4 (from Tables 6.1 and 6.2, column 4). The table compares those prior estimates to estimates on the combined *PIA*-*RAIS* subsample without (columns 2 and 5) and with log labor productivity as a regressor (columns 3 and 6). The export-status coefficient loses statistical significance in the reduced separation subsample of *PIA* manufacturers but does not change sign. Exporters exhibit significantly fewer accessions at the one-percent level; this reinforces the prior finding. Trade-variables, including Balassa comparative advantage, are not significant predictors of separations and accessions in the reduced subsample. Higher labor productivity itself, however, predicts significantly fewer accessions. This is consistent with the descriptive evidence (Table 7.2) that faster labor productivity growth at manufacturing firms correlates with slower-than-average workforce growth. Overall, the inclusion of log labor productivity in a smaller random sample of manufacturers overturns

Table 7.1: WORKER-EFFECT LOGIT ESTIMATION WITH LABOR PRODUCTIVITY

	Separations			Accessions		
	Full smpl.	*PIA* smpl.		Full smpl.	*PIA* smpl.	
	(1)	(2)	(3)	(4)	(5)	(6)
Balassa Comp. Adv.	.010	-.370	-.370	-.067	-.242	-.244
	(.045)	(.128)***	(.128)***	(.048)	(.120)**	(.120)**
Exporter Status	.293	.019	.019	-.438	-.287	-.281
	(.028)***	(.076)	(.076)	(.027)***	(.075)***	(.075)***
Product Market Tariff	-2.197	-.403	-.407	1.822	-.299	-.235
	(.469)***	(.976)	(.976)	(.498)***	(1.092)	(1.092)
Intm. Input Tariff	.265	.209	.223	-2.954	-.935	-1.114
	(.610)	(1.541)	(1.545)	(.750)***	(1.556)	(1.559)
Import Penetration	9.014	1.429	1.410	1.764	1.666	1.969
	(.463)***	(1.399)	(1.401)	(.665)***	(1.691)	(1.693)
Log Labor Productivity			.008			-.111
			(.051)			(.054)**
Obs.	145,408	40,335	40,335	112,974	20,191	20,191
Pseudo R^2	.144	.338	.338	.042	.092	.092

Source: Menezes-Filho and Muendler (2007). *RAIS* 1990-98 (1-percent random sample) linked to *PIA* 1990-98, male workers nationwide, 25 to 64 years old, with manufacturing job.
Note: Separations and accessions exclude transfers, deaths, and retirements. Reference observations are employments with no reported separation or accession in a given year. Sector information at subsector IBGE level. Estimates in column 1 and 4 repeat column 4 in Tables 6.1 and 6.2. Further regressors (not reported): Year indicators, sector indicators, sector, plant and worker covariates. Robust standard errors in parentheses: * significance at ten, ** five, *** one percent.

none of the previous results and reinforces several findings. Most import, faster productivity change depresses hiring rates, thus explaining lacking flows towards exporters and, by implication comparative advantage sectors, where productivity growth is relatively faster.

Labor and output reallocation. Much earlier research has emphasized the importance of market-share reallocations towards more productive firms after trade reform (e.g. Pavcnik 2003, Eslava et al. 2004). Those papers considered product-market shares in the reallocation analysis. None of that evidence is different for Brazil, where more productive firms also gain product-market shares. But, in the presence of endogenous 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 may remain 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 7.2 documents these distinct reallocation patterns in the product market and in the labor market. The table decomposes total factor productivity (columns 1-4) and labor productivity (columns 5-8) into the contributions of firm-level productivity and firm-level

Table 7.2: PRODUCTIVITY VARIATION ACROSS FIRMS AND OVER TIME

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

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

Source: Menezes-Filho and Muendler (2007). *PIA* firms 1986-98 (1991 missing).
Note: 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.
Note: Cross-sectional productivity decomposition as in Olley and Pakes (1996): $y_t = \bar{y}_t + \sum_i \overline{\Delta}\theta_{it}\overline{\Delta}y_{it}$, where y_t is weighted and $\bar{y}_t$ is unweighted mean log productivity and $\overline{\Delta}$ denotes deviations from cross-section means (rebased to unity in 1990). Annual productivity change correlation $\sum_{i\in C} \Delta\theta_{i,t}\Delta y_{i,t}$ (raw covariance) from Haltiwanger (1997) decomposition, where Δ denotes annual change (not rebased).

weights, where the weights are output in the case of total factor productivity and employment in the case of labor productivity (similar results also hold for total factor productivity and labor reallocation). The statistics are based on output and employment at formal-sector manufacturing firms. Following Olley and Pakes (1996), aggregate productivity in the cross section of firms (columns 1 and 5) is split into the unweighted mean productivity level (columns 2 and 6) and the covariance between deviations of the weights and productivities from annual means (columns 3 and 7). The relative log TFP change of 3.5 percent between 1990 and 1998 is modest (column 1).[17] Substantial capital accumulation contributes to the faster increase in log labor productivity by 7.3 percent between 1990 and 1998 (column 5). Alongside, Table 7.2 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.[18]

The decompositions in Table 7.2 show for the cross section of Brazilian manufacturers that firms with higher total factor productivity (TFP) do command larger output shares (column 3), and that TFP improvements among survivors are associated with gains in output shares (column 4). These facts are well known for Brazil and similar countries, but sometimes confounded with resource allocation. The cross-sectional covariance between labor productivity and employment shares, in fact, is considerably weaker (column 7) than between TFP and output shares (column 3). Most strikingly, firm-level labor productivity

[17]In Table 7.2, I 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.

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

advances are associated with reductions in employment shares (column 8).[19] So, firms with increasing productivity expand output shares but reduce employment.

Trade theories. While there is no explicit model for unilateral trade reforms and endogenous productivity change in response, recent trade theories investigate industry dynamics when trade costs drop worldwide in lock step and firms simultaneously engage in innovation and export-market participation.[20] Yeaple (2005) shows in a static model with ex ante identical firms and heterogeneous workers, whose skill is complementary to innovative technology, that the firms' binary choice of process innovation induces the sorting of more skilled workers to innovative firms, leading to firm heterogeneity ex post and to increased within-firm productivity in equilibrium. As multilateral trade costs drop, more firms in the differentiated-goods sector adopt innovative technology and raise their employment, hiring away the top-skilled workers from differentiated-goods producers with lower technology. Also considering ex ante identical firms, Ederington and McCalman (2008) allow for a continuous technology choice in a dynamic industry-equilibrium model and show that a drop in foreign trade costs raises the rate of technology adoption at exporters but delays it at non-exporters. Departing from ex ante heterogeneous firms, Costantini and Melitz (2008) reintroduce a stochastic productivity component from Hopenhayn (1992) into the Melitz (2003) model and allow firms to choose process innovation. In simulations of the dynamic industry equilibrium, an anticipated future reduction of multilateral trade costs leads firms to adopt innovation in advance, while waiting for export-market participation.

The mechanism by which productivity increases in these models is that globally reduced trade costs raise the returns from accessing the export market so that firms, which can both choose export-market participation and engage in innovation, adopt innovative technology because each activity raises the return to the other. Globally reduced trade costs are a carrot. Under a unilateral trade reform, in contrast, expected profits for domestic producers fall, potentially reducing incentives for innovation. So, unilateral trade reform is a stick. But it is a long-standing tenet in economics that product-market competition may discipline managers and workers by strengthening incentives in the respective principal-agent relationships. Stronger product-market competition may lead principals to become better informed (Hart 1983), induce managers to exert more effort to avert bankruptcy (Schmidt 1997), or lead surviving firms to strengthen incentives because induced exits of other firms raise profit opportunities (Raith 2003). This family of models, though never explicitly embedded in a trade context (or a general equilibrium context for that matter), offers a key explanation why firms may improve productivity in response to unilateral trade reform.

While endogenous productivity change in response to unilateral trade reform is absent from much of trade theory, classic and recent trade models nevertheless offer numerous predictions that are consistent with Brazil's experience. A particularly attractive model for

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

[20] As Arbache, Dickerson and Green (2004) have argued in the context of relative wage responses to trade before, standard trade theory ignores trade-induced technology adoption and implied relative labor-demand changes.

Table 7.3: WORKER-FIXED EFFECT LOGIT ESTIMATION WITH INTERACTIONS

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

Source: Menezes-Filho and Muendler (2007). *RAIS* 1990-98 (1-percent random sample), male workers nationwide, 25 to 64 years old, with manufacturing job.
Note: Separations and accessions exclude transfers, deaths, and retirements. Reference observations are employments with no reported separation or accession in a given year. Sector information at subsector IBGE level. Columns 1 and 4 repeat column 4 of Tables 6.1 and 6.2. Further regressors (not reported): Year indicators, sector, plant and worker covariates. Robust standard errors in parentheses: * significance at ten, ** five, *** one percent.

empirical work is the Bernard, Redding and Schott (2007) framework, which embeds heterogeneous firms in a classic trade model and derives 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 heterogeneity 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.[21]

An empirical investigation of the Bernard et al. (2007) model's labor-market predictions calls for the inclusion of higher-order interactions between trade reform, comparative advantage and exporting status. Table 7.3 compares previous separation and accessions estimates in columns 1 and 4 (from Tables 6.1 and 6.2, column 6) to regressions with interaction terms

[21] 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.

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). Interestingly, product tariff reductions depress accession rates most strongly in comparative-advantage industries, conditional on exporter presence. Similarly noteworthy is the fact that, 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. These findings are consistent with a main firm-level prediction of the Bernard et al. (2007) model: in comparative-advantage industries, existing jobs are destroyed less frequently at exporters. In contrast, the negative comparative advantage coefficient turns significant: employers in a comparative-advantage sector hire workers significantly less frequently. So, the classic-trade prediction that there is net job creation in comparative-advantage industries is statistically significantly challenged; comparative-advantage industries separate from significantly more workers and hire significantly fewer workers. This aspect of labor turnover, and the worker flows away from exporters, are arguably best explained by endogenous productivity change in response to trade reform.

8 Potential Implications for Adjustment Costs

The reported estimates owe their generality and robustness to a lean set of identifying assumptions. For the estimates in Sections 6 and 7, no structural assumption was needed other than that unobserved match-specific logistic shocks trigger separations or accessions beyond the observed variables. For the precise measurement of labor-market adjustment costs that are associated with trade reform, more explicit structural assumptions are required to model reallocation delays and failures in general equilibrium. Descriptive evidence in this section can nevertheless provide a benchmark assessment of the potential magnitude of adjustment costs.

The relevance of labor-market performance for aggregate outcomes. A chief concern of reallocation costs is with potentially idle labor: displaced workers who await formal-sector reallocation. They are not directly observable in formal-sector worker censuses. However, the Brazilian *RAIS* data like any formal-sector worker census record two margins that change the pool of prime-age male workers to be reallocated: separations from formal jobs fill the pool, and accessions into formal jobs empty the pool of workers to be reallocated.

Two important measures for the potential idleness of labor are the durations of successful reallocations within a given time period, such as twelve months following displacement, and the rate of failed reallocations within the time horizon. Numerous economic causes can be responsible for changes to the the durations of successful reallocation and the rate of failed reallocations. Without any reference to possible causes, Table 8.1 shows the changes to reallocation durations and failures over time. The table documents that the share of displaced workers without reallocation for a year almost doubles from 25 to 46 percent between 1986 and 1998. There is some variation in the failure rate across skill groups within any given

Table 8.1: Labor Market Performance and Economic Outcomes

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

Source: Menezes-Filho and Muendler (2007).
Note: *RAIS* 1986-1999 (1-percent random sample), male workers nationwide, 25 to 64 years old, displaced from a formal-sector job; not rehired into a formal-sector job within 12 months (*upper panel*) or rehired into a formal-sector job within 12 months (*lower panel*). *PME* 1986-1999, share of idle workers (unemployed or withdrawn from labor force), and *Banco Central do Brasil*, GDP. Young workers have ten or less years of potential labor force experience, and college-educated workers have some college education. Foregone GDP is the unrealized wage bill, measured as the product of the observed change over 1990 times the number of newly displaced workers during the year times their wage upon displacement. Idle labor is defined as the share of displaced workers with transitions to unemployment or out of the labor force.

year: young and college-educated workers' reallocations fail less frequently than average. Time variation, however, dwarfs the skill-group differences. A similar pattern applies to durations of successful reallocations in the lower panel of Table 8.1. The relatively minor cross-sectional differences between skill groups, compared to major time variation, suggests that studying macroeconomic sources of variation in labor-market performance promises to uncover first-order changes in labor-market outcomes.

Idle resources in the labor market are a foremost component of Brazil's aggregate performance. For reallocation failures in the upper panel of Table 8.1, the foregone share of GDP is calculated as the unrealized wage bill that the additional failures after 1990 imply, given a displaced worker's last wage. This computation only considers the share of displaced formal-sector workers as idle who typically become unemployed or move out of the labor force—a 36 percent share on average in *PME* 1990-98. So, the assumption is that the remaining 64 percent of displaced workers immediately take up an informal job or self employment and fully retain their pre-displacement earnings. This makes the estimates of foregone GDP conservative. The magnitudes are nevertheless striking. The unrealized wages implied by additional reallocation failures after 1990 amount to 1.4 percent of foregone GDP in 1992 and 2.4 percent in 1998. The increased duration of successful reallocations in the lower panel of Table 8.1 implies another half percent of foregone GDP in 1992 and .6 percent in 1998. This brings the total foregone wage bill to almost 2 percent of GDP in 1992, to almost 1 percent in 1994 (a year with strong GDP growth), and to 3 percent in 1998. The estimates

Table 8.2: TRADE EXPOSURE AND PREDICTED LABOR MARKET OUTCOMES

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

Source: Menezes-Filho and Muendler (2007). *RAIS* 1990-98, male workers nationwide, 25 to 64 years old, with manufacturing job (estimation samples from Tables 6.1 and 6.2).
Note: Sector information at subsector IBGE level. Predicted changes in separation and accession rates based on marginal effects implied by column (6) estimates in Tables 6.1 and 6.2 ($\hat{P}(1-\hat{P})$ is .170 for separations and .174 for accessions).

are conservative because the calculations only consider two out of five employees, prime-age male workers, and because it is assumed that displaced workers who become informal or self-employed retain the full pre-displacement earnings immediately after displacement.

The relevance of trade for labor-market performance. To gain a sense of how important trade is for labor-market performance in Brazil's manufacturing sector, one can use the changes in import penetration rates and tariffs since 1990 to predict changes in separation and accession rates, based on conditional logit estimates from Tables 6.1 and 6.2 (columns 5) in Section 6. Import penetration more than doubles between 1990 and 1998, while product-market and input tariffs drop by more than half. As Table 8.2 reports, additional import penetration in 1998, beyond its 1990 level, predicts a 6-percent increase in the separation rate by 1998. Tariff reductions below 1990 levels predict an 8-percent increase in the separation rate. These are salient magnitudes compared to the mean separation rate of .282 over the 1990s (Table 3.2). The manufacturing sector employs roughly one in four male workers in the national labor force, so these predictions matter for the aggregate. At the accession margin, additional import penetration predicts a 4-percent reduction in hiring rates, whereas the ambiguous effect of product and input tariffs on accession rates partly counteracts the prediction. Overall, the magnitudes suggest that trade is a potentially important factor that affects national labor-market performance.

9 Conclusion

Brazil's labor-market adjustment after large-scale trade reform in the early 1990s offers important insights into prospective reallocation shifts and associated adjustment costs from rising reallocation durations and failure rates. A conventional labor-demand decomposition documents two salient workforce changeovers. Within the traded-goods sector, there is a marked occupation downgrading and a simultaneous education upgrading, by which employers fill expanding low-skill intensive occupations with increasingly educated jobholders. Between sectors, there is a labor demand shift towards the least and the most skilled, which can be traced back to relatively weaker declines of traded-goods industries that intensely use low-skilled labor and to relatively stronger expansions of nontraded-output industries that intensely use high-skilled labor. These observations are broadly consistent with predictions of classic trade theory for a low-skill abundant economy.

Actual worker flows, however, reveal a much more nuanced picture. Rich linked employer-employee data show that workforce changeovers are neither achieved through worker reassignments to new tasks within employers nor are they brought about by reallocations across employers and traded-goods industries. Instead, trade-exposed industries shrink their workforces by dismissing less-schooled workers more frequently than more-schooled workers. Most displaced workers shift to nontraded-output industries or out of recorded employment. Brazil's trade liberalization triggers worker displacements particularly from protected industries, as trade theory predicts and welcomes. But neither comparative-advantage industries nor exporters absorb trade-displaced workers. To the contrary, comparative-advantage industries and exporters displace significantly more workers and hire fewer workers than the average employer, and resource reallocation appears to remain incomplete for years. These patterns are best explained by relatively fast labor productivity increases at exporters and in comparative advantage industries. Employers in those activities raise productivity in response to heightened competition, and they do so faster than nonexporters and firms in disadvantage industries because expected exporting activity increases the return to innovation. As a result, product market shares shift to more productive firms. Product market-shares grow less than proportional with productivity, however, so that trade-induced productivity growth leads to labor savings at exporters and in comparative advantage industries.

The labor-market evidence for Brazil also offers a novel explanation why pro-competitive reforms can be associated with strong efficiency gains at the employer level but not in the aggregate, where idle resources result. Conservatively measured, the foregone wage bill from the increase in reallocation durations and failures after 1990 amounts to between one and three percent of GDP. Of course, these computations are by no means a substitute to rigorous adjustment-cost modelling and measurement. The magnitudes are nevertheless indicative of potentially important adjustment costs that take on a magnitude of foregone GDP similar to the business cycle. The magnitude of adjustment costs is, however, not likely close to the repeated static gains from trade, which are typically thought to amount to welfare gains in the order of large fractions of GDP. A promising path for future research is the rigorous theoretical modelling and empirical measurement of adjustment cost in the labor market in response to trade integration.

Appendix

A Linked employer-employee data

The main data source underlying statistics in this summary note is linked employer-employee data. Brazilian law requires every Brazilian plant to submit detailed annual reports with individual information on its employees to the ministry of labor (*Ministério de Trabalho*, MTE). The collection of the reports is called *Relação Anual de Informações Sociais*, or *RAIS*, and typically concluded at the parent firm by late February or early March for the preceding year of observation. *RAIS* primarily provides information to a federal wage supplement program (*Abono Salarial*), by which every worker with formal employment during the calendar year receives the equivalent of a monthly minimum wage. *RAIS* records are then shared across government agencies. An employer's failure to report complete workforce information can result in fines proportional to the workforce size; but fines are seldom issued. A strong incentive for compliance is that workers' benefits depend on *RAIS* so that workers follow up on their records. The payment of the worker's annual public wage supplement is exclusively based on *RAIS* records. The ministry of labor estimates that currently 97 percent of all formally employed workers in Brazil are covered in *RAIS*, and that coverage exceeded 90 percent throughout the 1990s.

Observation screening. In *RAIS*, workers are identified by an individual-specific PIS (*Programa de Integração Social)* number that is similar to a social security number in the United States (but the PIS number is not used for identification purposes other than the administration of the wage supplement program *Abono Salarial*). A given plant may report the same PIS number multiple times within a single year in order to help the worker withdraw deposits from the worker's severance pay savings account (*Fundo de Garantia do Tempo de Serviço*, *FGTS*) through spurious layoffs and rehires. Bad compliance may cause certain PIS numbers to be recorded incorrectly or repeatedly. To handle these issues, I screen *RAIS* in two steps. (1) Observations with PIS numbers shorter than 11 digits are removed. These may correspond to informal (undocumented) workers or measurement error from faulty bookkeeping. (2) For several separation statistics, I 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, I only keep the observation with the highest average monthly wage level (in cases of wage ties, I 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, age in years is categorized into those eight age ranges also for 1994-2001. I 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. I 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	Years
1.	Illiterate, or Primary or Middle School Educated	1-5	0-8
2.	Some High School or High School Graduate	6-7	8-12
3.	Some College	8	12+
4.	College Graduate	9	16+

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.[22] I map ISCO-88 categories to *RAIS* occupations as follows:

	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

Employment. Table A.1 shows the employment allocation by industry in the universe of *RAIS* workers in 1986, 1990 and 1997.

[22] See documentation at URL *econ.ucsd.edu/muendler/brazil*.

Table A.1: EMPLOYMENT ALLOCATION BY SUBSECTOR

Sector and subsector IBGE		Employment share 1986 (1)	1990 (2)	1997 (3)
Primary				
1	Mining and quarrying	.007	.006	.004
25	Agriculture, farming, hunting, forestry and fishing	.015	.016	.041
Manufacturing				
2	Manufacture of non-metallic mineral products	.016	.013	.011
3	Manufacture of metallic products	.030	.024	.021
4	Manufacture of machinery, equipment and instruments	.020	.016	.011
5	Manufacture of electrical and telecommunications equipment	.016	.014	.008
6	Manufacture of transport equipment	.019	.016	.013
7	Manufacture of wood products and furniture	.019	.015	.015
8	Manufacture of paper and paperboard, and publishing	.014	.014	.013
9	Manufacture of rubber, tobacco, leather, and products n.e.c.	.019	.016	.009
10	Manufacture of chemical and pharmaceutical products	.024	.022	.020
11	Manufacture of apparel and textiles	.042	.035	.026
12	Manufacture of footwear	.012	.010	.008
13	Manufacture of food, beverages, and ethyl alcohol	.040	.039	.041
Commerce				
16	Retail trade	.106	.103	.127
17	Wholesale trade	.024	.025	.027
Services				
18	Financial intermediation and insurance	.038	.034	.025
19	Real estate and business services	.074	.073	.079
20	Transport, storage and telecommunications	.050	.044	.057
21	Hotels and restaurants, repair and maintenance services	.101	.101	.084
22	Medical, dental and veterinary services	.014	.017	.039
23	Education	.008	.009	.036
Other				
14	Electricity, gas and water supply	.013	.014	.014
15	Construction	.045	.041	.049
24	Public administration and social services	.209	.206	.224
26	Activities n.e.c.	.025	.077	.001
Total employment (thousands of workers)		22,164	23,174	24,104

Source: *RAIS* 1986, 1990 and 1997, universe of workers.
Note: Employment on Dec 31. Slight differences to Table 3.1 are due to random sampling errors.

B Additional data sources

Throughout this summary note, I draw on several additional data sources. The Brazilian monthly employment survey *PME* (*Pesquisa Mensal de Emprego*) is conducted by Brazil's statistical bureau IBGE, using a rotating panel. *PME* follows households for 16 months, with an eight-month interval after the fourth interview.[23] Changes to the sample design adversely affect worker panels starting in odd years. So, I use only individuals whose first survey occurs in 1986, 1988, 1990, 1992, 1994, 1996 or 1998. As with *RAIS*, I restrict our sample to prime-age male workers. In the survey, individuals without employment are considered unemployed if they report active search for work during the week prior to the interview, and are considered out of the workforce otherwise. Household members who work for their own account but do not employ others are considered self-employed. I exclude individuals who become employers.

For descriptive statistics in Table 7.2 and robustness checks in Table 7.1, I 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. I first obtain log TFP measures from Olley and Pakes (1996) estimation at the *Nível 50* sector level under a Cobb-Douglas specification (Muendler 2004). I then convert log TFP to log labor productivity by adding the production-coefficient weighted effects of capital accumulation and intermediate input use. Labor productivity is denominated in BRL-deflated USD-1994 output equivalents per worker. IBGE's publication rules allow data from *PIA* to be withdrawn in the form of tabulations with at least three firms per entry. I 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. I 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, I calculate mean log productivity but retain the firm identifiers behind the combination—permitting the linking to *RAIS*.

I use data on *ad valorem* tariffs by sector and year from Kume, Piani and Souza (2003). The tariffs are the legally stipulated nominal rates for Brazil's trade partners with no preferential trade agreement, and not weighted by source country. I combine these tariff series with economy-wide input-output matrices from IBGE to arrive at intermediate input tariff measures by sector and year. I calculate the intermediate-input tariff as the weighted arithmetic average of the product-market tariffs, using sector-specific shares of inputs for the input-output matrix as weights. I use Ramos and Zonenschain (2000) national accounting data to calculate market penetration with foreign imports. Arguably, domestic firms find the absorption market, corresponding to output less net exports, the relevant domestic environment in which they compete. I define the effective rate of market penetration as imports per absorption. Foreign direct investment (FDI) and annual GDP data are from the Brazilian

[23] 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 $m+15$.

central bank.

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

C Wage structure in manufacturing

Table C.1 presents Mincer (1974) regressions of the log wage on individual compensation components. Following Abowd, Kramarz, Margolis and Troske (2001), individual compensation in a given year is given by

$$\ln w_i = x_i\beta + \psi_{J(i)} + \varepsilon_i, \tag{C-1}$$

where w_i is worker i's annual wage, x_i is a vector of observable worker characteristics including gender, experience, education and occupation, β is a vector of parameters to be estimated, $\psi_{J(i)}$ is an plant effect ($j = J(i)$ being the plant that employs worker i), and ε_i is an error term. The plant effect combines a pure plant effect with the plant average of pure worker effects:

$$\psi_j = \phi_j + \overline{\alpha}_j, \tag{C-2}$$

where ϕ_j is the pure plant effect and $\overline{\alpha}_j$ is the average of pure worker effects α_i over workers employed at plant j. The plant effect controls for unobservable worker and plant characteristics. Abowd and Kramarz (1999) show that omitting this effect leads to bias in the estimation of β in general.

Regressors are potential worker experience and indicator variables for gender, education and occupation as measures of individual characteristics. Quadratic, cubic and quartic terms for potential experience are included. Gender is interacted with all other variables. Table C.1 presents results for the manufacturing sector in São Paulo state in 1990 and 1997. Comparable estimates for manufacturing workers in France in 1992 and the U.S. in 1990, drawn from Abowd et al. (2001), are also reported.[24]

[24]Data for France derive from the *Enquête sur la Structure des Salaires* (*ESS*), which samples responses to an annual administrative census of business enterprises. Data for the U.S. derive from the *Worker-Establishment Characteristic Database* (*WECD*), which links individual census responses to manufacturing plants surveyed in the *Longitudinal Research Database* (*LRD*). See Abowd et al. (2001) for further details.

Table C.1: MANUFACTURING WAGES IN BRAZIL, FRANCE AND THE U.S.

	Brazil 1990	Brazil 1997	France 1992	U.S. 1990
	(1)	(2)	(3)	(4)
Primary School Education (or less)	-1.075	-1.000	-.338	-.526
	(.002)	(.002)	(.009)	(.008)
Some High School Education	-.923	-.881	-.256	-.404
	(.002)	(.002)	(.009)	(.007)
Some College Education	-.339	-.316	-.200	-.334
	(.003)	(.003)	(.009)	(.007)
College Graduate			-.064	-.123
			(.016)	(.007)
Professional or Managerial Occupation	.856	.912	.760	.359
	(.002)	(.002)	(.009)	(.004)
Technical or Supervisory Occupation	.600	.632	.401	.206
	(.002)	(.002)	(.007)	(.004)
Other White Collar Occupation	.262	.249	.169	-.039
	(.002)	(.002)	(.011)	(.005)
Skill Intensive Blue Collar Occupation	.239	.225	.155	.083
	(.001)	(.001)	(.007)	(.003)
Potential Labor Force Experience	.095	.082	.069	.083
	(.0005)	(.0007)	(.003)	(.002)
Quadratic Experience Term	-.003	-.003	-.004	-.003
	(.00005)	(.00007)	(.0002)	(.0001)
Cubic Experience Term	.00005	.00008	.0001	.00007
	(2.29e-06)	(2.86e-06)	(1.00e-05)	
Quartic Experience Term	-3.01e-07	-7.64e-07	-1.20e-06	-4.70e-07
	(3.24e-08)	(3.89e-08)	(1.00e-07)	(3.00e-08)
Female	.060	.070	.052	-.078
	(.005)	(.006)	(.024)	(.019)
Female × Primary School Education (or less)	.106	.051	*-.0006*	.041
	(.004)	(.004)	(.021)	(.016)
Female × Some High School Education	-.016	-.058	*-.016*	*-.009*
	(.004)	(.004)	(.021)	(.015)
Female × Some College Education	.018	*-.005*	*.025*	*-.019*
	(.005)	(.005)	(.021)	(.015)
Female × College Graduate			-.062	*-.022*
			(.029)	(.015)
Female × Professional or Managerial Occupation	-.101	-.058	-.049	-.086
	(.004)	(.005)	(.016)	(.007)
Female × Technical or Supervisory Occupation	-.173	-.250	*-.006*	.037
	(.003)	(.004)	(.011)	(.008)
Female × Other White Collar Occupation	.088	.071	.033	.046
	(.003)	(.003)	(.013)	(.006)
Female × Skill Intensive Blue Collar Occupation	-.208	-.167	-.045	-.043
	(.002)	(.003)	(.010)	(.008)
Female × Potential Labor Force Experience	-.056	-.036	-.047	-.016
	(.0008)	(.001)	(.004)	(.003)
Female × Quadratic Experience Term	.002	.002	.004	*.0003*
	(.0001)	(.0001)	(.0003)	(.0002)
Female × Cubic Experience Term	-.00006	-.00005	-.0001	.00000
	(4.35e-06)	(5.63e-06)	(1.00e-05)	
Female × Quartic Experience Term	7.06e-07	5.40e-07	1.20e-06	*1.80e-08*
	(6.32e-08)	(7.78e-08)	(1.10e-07)	(4.00e-08)
R^2 (within)	.508	.468	.817	.617
Residual degrees of freedom	2,326,428	1,828,049	23,920	148,992

Sources: Menezes-Filho et al. (2008). *RAIS* São Paulo state manufacturing 1990 and 1997 (prime age workers in their highest-paying job), Abowd et al. (2001) for France and the U.S., controlling for plant fixed effects.
Note: Estimates for Brazil relative to college graduates, for France and the U.S. relative to workers with post-graduate degree. Robust standard errors in parentheses: * significance at ten, ** five, *** one percent. .

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