Labor Reallocation in Response to Trade Reform*

Naércio Aquino Menezes Filho‡

IBMEC São Paulo and Universidade de São Paulo

Marc-Andreas Muendler¶

University of California, San Diego and CESifo

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Abstract

Tracking individual workers across sectors and firms after Brazil's trade liberalization in 1990 shows that foreign import penetration triggers worker displacements but that neither comparative-advantage sectors nor exporters absorb displaced workers for years. Displacements are significantly more and accessions significantly less frequent in comparative-advantage sectors and at exporters. Heightened import penetration is consequently associated with significantly more frequent transitions to informal work status and unemployment, and longer durations and more frequent failures of formal-job reallocations. So, the output reallocation to more productive firms after trade reform is not accompanied by similar labor reallocation. The sector and firm affiliation of high-turnover workers is not random. JEL F14, J23, J63

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‡naercioamf@isp.edu.br (*www.econ.fea.usp.edu/naercio*)

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

1 Introduction

Latin America's economic growth in the 1990s is widely perceived as disappointing in the wake of pro-competitive reforms that were expected to boost aggregate performance. Whereas several empirical studies at the level of plants and firms present evidence of considerable productivity advances, both within firms and through output reallocations between them, aggregate measures of GDP per capita do not exhibit strong growth.[1] We document at the level of plants and their individual workers for Brazil that a reason for the weak aggregate growth performance may be imperfectly functioning labor markets.

When factor productivity improves, a favorable reallocation of output shares to more productive firms need not be, and is not in Brazil, accompanied by labor reallocation to the more productive firms. While Brazilian manufacturing firms with productivity gains command larger output shares on average, these manufacturing firms reduce their employment share in the workforce on average. We document in worker level regressions that Brazil's dismantling of trade barriers in the early 1990s significantly increases worker transitions from formal jobs into informality and unemployment, and significantly reduces the transition from informality back into formal jobs. We show, moreover, that workers in sectors with a comparative advantage and workers at exporting firms suffer significantly more frequent displacements. Worse, worker accessions to jobs in comparative advantage sectors and at exporters are also less likely. Consequently, sectors and firms that are commonly expected to expand after trade reform fail to absorb the displaced workers for a period of at least several years—contrary to the factor-market premises of classic and firm-level trade theories.

We focus our analysis on male workers with ages between 25 and 64, whose labor supply decision is arguably less sensitive to economic change after first entry into the labor force. We use a metropolitan household survey (*PME*) to obtain direct information on prime-age male workers with no formal-sector employment—the focal group of workers to-be-reallocated. To identify employers and their export status, to track migrants, and to condition on worker-fixed

[1] Bosworth and Collins (2003) report for Latin America an average annual output-per-worker growth rate of .9 percent, of which they attribute .4 percent to total factor productivity change. Among the micro-level studies, Eslava, Haltiwanger, Kugler and Kugler (2004) report TFP estimates that imply an annual TFP growth rate of 3.5 percent between 1990 and 1998 at Colombian manufacturing plants; Schor (2004) reports TFP estimates sector by sector that imply an average annual TFP growth rate of 1.9 percent between 1990 and 1998 at Brazilian manufacturing firms.

effects known to workers and employers but otherwise unobserved, we use worker panels from the national linked employer-employee records *RAIS*. We find the sector and firm affiliation of high- and low-turnover workers to be non-random so that worker characteristics matter for trade-related covariates of separation, accession, and transition probabilities.

Though hard to reconcile with classic trade theory and frictionless firm-level models, our findings are in line with alternative explanations. Aspects of Brazil's experience could be interpreted as consistent with predictions of recent trade models that make factor-market institutions a source of comparative advantage and find that countries with less rigid factor markets tend to export products from industries with high factor turnover (Saint Paul 1997, Davidson, Martin and Matusz 1999, Cunat and Melitz 2006). Brazil's comparative-advantage sectors exhibit more labor turnover: higher worker separation rates and unconditionally higher worker accession rates. If one is willing to accept the assertion that Brazil's large informal sector and elevated annual separation and accession rates in the formal sector make its labor market relatively flexible, then this aspect is consistent with factor-market-institutions models.

The subdued expansion of comparative advantage sectors, however, is not necessarily compatible with that set of explanations either. Rigid real wages, which increase throughout the 1990s in Brazil, are a known cause for hampered reallocation in trade models (Brecher 1974). Simultaneous productivity change and, by extension, the use of substitute factors such as foreign capital and intermediates can be shown to reduce employment in general equilibrium, depending on consumer-demand elasticities (e.g. Obstfeld and Rogoff 1996, 4.3.2). However, the prolonged time horizon of Brazil's labor market adjustment—with 1.5 times longer reallocation durations and more frequent reallocation failures over a period of at least eight years after trade reform—speaks to considerable adjustment costs.

We proceed as follows. Section 2 elaborates the main facts about market adjustment after trade reform in Brazil, relating the facts to prior evidence and trade theories. After a discussion of the data (Section 3) and estimation models (Section 4), we show in worker-level regressions in Section 5 that Brazil's market adjustment is indeed statistically significantly related to its changing trade regime. Firms simultaneously determine export status and labor turnover, however, and Brazil's tariff reductions and subsequent import penetration are targeted at low-efficiency sectors, while the economy undergoes a progressive exchange rate revaluation, privatization, foreign investment inflows, some adjustment to preceding labor-market reforms, and service-job outsourcing. Section 6 presents a series of according statistical treatments to corroborate our hypothesis that trade reform does trigger the market adjustments and that comparative-advantage sectors and exporters fail to absorb displaced labor. Section 7 concludes.

2 New Facts, Earlier Findings, and Theories

Reallocation of labor is the reassignment of workers to jobs across plants and sectors. In the presence of plant exit and productivity change, labor reallocation is distinct from the reallocation of output shares. Take firm exit as an example. If firms exit but displaced workers are not reabsorbed at surviving or new firms in the formal sector, labor reallocation remains incomplete while output shares are being shifted to survivors and entrants. So, a positive correlation of output shares with productivity is no evidence that factor reallocation succeeds.

Table 1 decomposes total factor productivity (columns 1-4) and labor productivity (columns 5-8) into the contributions of firm-level productivity and firm-level weights, where the weights are output in the case of total factor productivity and employment in the case of labor productivity. 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). Alongside, Table 1 reports the raw covariance of year-over-year changes in weights and productivities at surviving firms (columns 4 and 8)—a term in the Haltiwanger (1997) decomposition over time. The overall TFP gains between 1990 and 1998 are modest in the *PIA* manufacturing firm sample (column 1).[2] Due to substantial capital accumulation at the firms, however, labor productivity improves faster (column 5).

The decompositions in Table 1 confirm for the cross section of Brazilian manufacturers that firms with higher total factor productivity (TFP) do command larger output shares (column 3), and that TFP improvements among survivors are associated with gains in output shares (column 4). The cross-sectional covariance between labor productivity and employment shares, however, is considerably weaker (column 7) than between TFP and output shares (column 3).

[2] In Table 1 we divide aggregate log productivity levels by the aggregate 1990 log level. This results in an apparent TFP increase of only 3.5 percent between 1990 and 1998, whereas rebasing to 1986 at the firm level in Muendler (2004) yields a 4.7 percent increase between 1990 and 1998.

Table 1: PRODUCTIVITY ACROSS FIRMS AND OVER TIME

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

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

Source: *PIA* firms 1986-98 (1991 missing); log total factor productivity from Muendler (2004) based on Olley and Pakes (1996) estimation (at *Nível 50*), inferring labor productivity at changing capital stocks. 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).

Most strikingly, firm-level labor productivity advances are associated with reductions in employment shares (column 8).[3] So, more productive firms command increasing output shares while reducing employment.

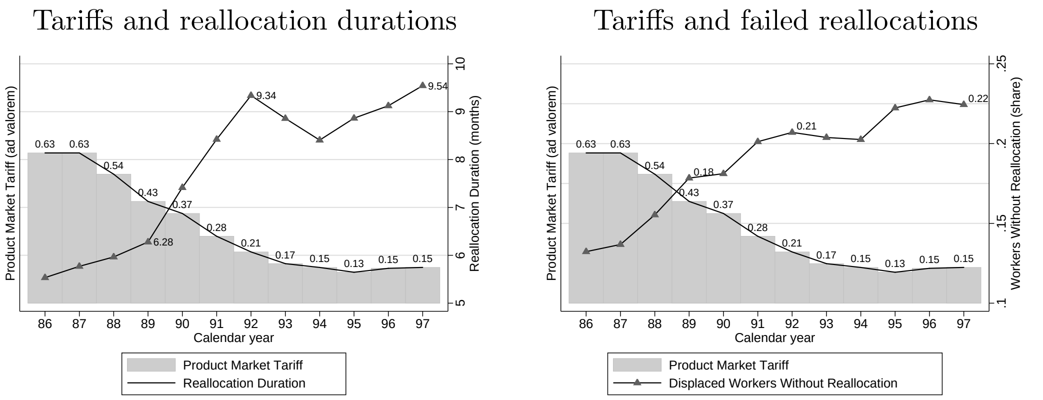
Trade reform and labor-market performance. In 1988, after decades of import substitution in Brazil, the Sarney government formally begins to reduce nominal tariffs but leaves binding non-tariff barriers largely untouched. In 1990, the newly elect president Collor de Melo picks a far-reaching reform plan from several competing ministerial proposals, issues a presidential directive to eliminate non-tariff barriers and special import regimes on his first day in office, and presents a detailed schedule for tariff reductions to be completed by 1994. The sudden enactment and the far-reaching changes to the trade regime take Brazilian businesses by surprise. The government's declared objectives for dismantling trade barriers are, first, to instill competition in inefficient sectors and, second, to discipline concentrated industries in their pricing power so that hyper-inflation can be fought effectively. As a consequence, and contrary to common political-economy outcomes, mostly sectors with sluggish efficiency performance are tar-

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

geted with reduced tariffs. (Our regression analysis will control for the induced simultaneity in tariff levels and sector performance.) The trade liberalization programme is completed in less than three years by July 1993. Negotiations for the Southern Cone (*Mercosur*) trade agreement with Argentina (Brazil's second largest trading partner after the U.S. during the 1990s), Paraguay and Uruguay are concluded by December 1994. Tariff resettings under *Mercosur* take effect in January 1995 and reverse liberalization efforts in select sectors, but for a majority of industries Brazil's trade barriers remain at their 1993 levels.

Brazil's unilateral trade reform of 1990 removes price distortions at the border. This tends to increase the volume of trade, all else equal. In fact, Brazil's exports increase from 8.2 to 10.9 percent of GDP between 1990 and 1992 in the immediate aftermath of trade reform, but drop back to 7.3 percent by 1998 under the progressive overvaluation of Brazil's Real until January 1999. (Our regression analysis condition on real exchange rate effects.) Heckscher-Ohlin-Samuelson and Ricardo-Viner trade theory predicts that the mobile labor force is reallocated to sectors with a comparative advantage, all else equal. Multilateral tariff reductions among Brazil's Southern Cone trading partners benefit Brazil's exporters, and plant-level trade theory (Bernard, Eaton, Jensen and Kortum 2003, Melitz 2003) predicts that labor is reallocated to expanding exporters, given plant productivity. Classic and firm-level trade theories alike show the desirability of worker reallocations: gains from specialization, beyond gains from exchange, accrue when factors are reassigned to activities with a comparative advantage. Neither Brazil's comparative-advantage sectors nor Brazil's exporters, however, provide the necessary reabsorption of displaced workers for these gains from specialization to materialize during the 1990s.

There are several criteria to assess the ensuing labor reallocation. Figure 1 depicts declining tariff levels in Brazil alongside two measures of the reallocation success. The left graph shows the mean duration in months until displaced workers are reabsorbed into a formal-sector job. Whereas it takes workers who suffer displacement in 1989 6.3 months to find another formal-sector job, after trade reform the mean duration increases to 9.3 months and remains elevated. Simultaneously, the share of displaced workers who fail to find another formal-sector job within four years increases from 18 percent in 1989 to more than 21 in 1992 and remains elevated too. There is regional heterogeneity behind the nationwide statistics but overall trends are similar across regions: the share of displaced workers with successful reallocations in the first year of the four-year horizon drops from 83 to 75 percent nationwide and from 97 to 94 percent in metropolitan areas; conversely, the share of workers who take one to three years



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

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

to be successfully reallocated within four years increases from 15 to 20 percent nationwide, and from 3 to 5 percent in metropolitan areas.

Additional large-scale reforms affect Brazilian labor markets during the same period. First, Brazil's new constitution, enacted in 1988, alters labor regulations: among other reforms, it reduces the work week from 48 to 44 hours, raises extra vacation pay to about a third of the monthly wage, and increases dismissal compensation fourfold. Second, a pro-competitive privatization programme is launched in 1995. (We will account for both changes in regression analysis and provide evidence that trade reform triggers displacements and hampers reallocation beyond the effects of constitutional change and privatization.) In the time series, a factor behind the lacking improvement in reallocationss after 1994 may be privatization. Similarly, a reason for the deterioration in reallocations between 1988 and 1990 could be changing labor-market institutions. Paes de Barros and Corseuil (2004) find no evidence, however, that the constitutional change in 1988 had a significant effect on labor demand. Whereas separation rates for briefly employed workers drop slightly, they increase slightly for workers with longer plant tenure, leaving no detectable net effect on employment.

Work status. For worker welfare, an important aspect of labor reallocation is the transition between work status categories. The slower and now frequently

Table 2: Employment by Sector and Formality Status

	Primary	Manuf.	Comm.	Services	Other	*Total*[a]
	(1)	(2)	(3)	(4)	(5)	(6)
			Allocation nationwide (*RAIS*)			
1990	.029	.263	.111	.284	.314	10.763
1998	.064	.207	.134	.308	.286	11.640
			Allocation in metropolitan areas (*RAIS*)			
1990	.015	.270	.104	.309	.302	5.965
1998	.023	.198	.125	.369	.285	6.057
			Informality share in metropolitan areas (*PME*)			
1990	.159	.063	.109	.117	.298	
1998	.232	.120	.154	.169	.341	

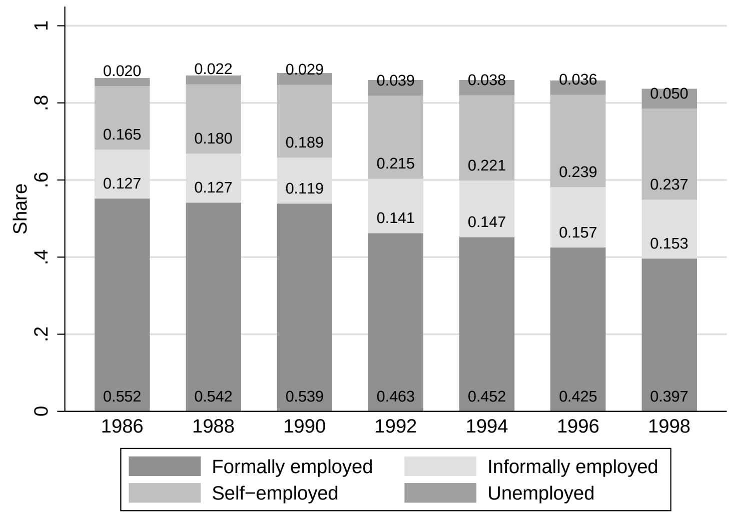
[a]Total employment (million workers), scaled to population equivalent.

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

failing reallocation in the formal sector is accompanied by a relative contraction of the manufacturing sector and an increase in informal work status across all sectors of the Brazilian economy. Between 1990 and 1998, the share of prime-age male workers with manufacturing employment drops from 26 to 21 percent in the national formal labor force (Table 2), and from 27 to 20 percent in the metropolitan formal labor force. The increase in informality is most pronounced in manufacturing, where the frequency of informal work status almost doubles from slightly more than six percent to twelve percent in the metropolitan labor force (for which we have data on work status).[4]

Figure 2 shows changes to work status. Whereas 53 percent of metropolitan males are formally employed in 1990, only 40 percent hold a formal job by 1998. Most male household members who drop from the formal workforce seek self-employment, which increases from 19 to 24 percent of the labor force between

[4]Unconditionally, wages in informal jobs in Brazil are below those in comparable formal jobs but, when controlling for worker characteristics and their selectivity into informal work states, informal jobs pay a wage premium that partly compensates for the foregone benefits from formal employment protection and old-age security (Menezes Filho, Mendes and de Almeida 2004).



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

Figure 2: **Work status in metropolitan areas**

1990 and 1998; many male workers change to informal employment, which rises from below 12 to above 15 percent; and several male household members suffer unemployment, which increases from 3 to 5 percent. The share of male household members who declare to have withdrawn from the labor force rises from less than 13 to more than 16 percent.

Labor reallocation across sectors and firms. For economic gains from specialization, a crucial aspect of labor reallocation is the transition of workers between sectors and types of firms. Linked employer-employee data allow us to track workers between jobs over time, and provide a detailed picture of transitions between sectors and firms.

For a panel of sectors in six Latin American countries, Haltiwanger, Kugler, Kugler, Micco and Pagés (2004) find that tariff reductions are associated with heightened within-sector churning—the excess gross job creation and destruction beyond observed net employment changes—, and with net employment reductions at the sector level.[5] At the individual country level, Revenga

[5]In a related approach, Roberts (1996) does not detect a clear effect of trade exposure

(1997) documents that Mexico's tariff cuts are associated with reduced firm-level employment in manufacturing. For Brazil, Ribeiro, Corseuil, Santos, Furtado, Amorim, Servo and Souza (2004) compute job creation and destruction rates at the sector level and report that lower tariff barriers are associated with increased job destruction, but with no significant effect on job creation. These employment contractions in formerly protected sectors are both expected and desirable: trade theory welcomes factor displacements from activities with a comparative disadvantage, which used to be protected under import substitution. But does a successful reallocation to comparative-advantage activities ensue? It takes worker panel data to scrutinize the cross-sectoral reabsorption after displacement.

Table 3 reports transitions of displaced formal-sector workers to other formal-sector jobs at the annual horizon (columns 1-6) between 1986 and 2001, and the share of displaced workers with no observed formal-sector rehiring within a year (column 7). Retained workers do not enter the statistics. Agricultural, mining and manufacturing plants are grouped into their sector's comparative advantage quintiles at the subsector *IBGE* level (comparable to the *NAICS* three-digit level). We define a sector's comparative advantage as its share in Brazilian exports relative to the sector's world export share (Balassa 1965).[6] Output in all other sectors is considered nontraded.

The majority of successful reallocations of workers within traded goods sectors is to plants in the same comparative-advantage quintile: transition rates along the diagonal in the five traded sectors exceed those off the diagonal (column 1-5). Transitions to sectors with similar trade exposure occur more frequently than to dissimilar sectors: off-diagonal entries are small, especially for accession sectors whose comparative advantage rank is two or more quintiles away from the separation sector. These facts suggest that traded-goods sectors with different degrees of comparative advantage are little permeable to labor reallocation. Classic trade theory predicts, to the contrary, that factors

on employment changes in Chile and Colombia, once sector characteristics are taken into account. Neither do Davis, Haltiwanger and Schuh (1996) identify a clear effect of trade on factor reallocation using U.S. data, nor do Wacziarg and Wallack (2004) detect a statistically systematic pattern of labor reallocation across formal manufacturing sectors in a cross-country study around trade-liberalization periods. Studies considering exchange rate effects (Klein, Schuh and Triest 2003, Gourinchas 1999), however, do find systematic effects on employment flows.

[6]For details on data construction, see Section 3 (and Table 17, p. 42, with manufacturing quintiles). Results with the *CNAE* sector classification are (comparable to the *NAICS* four-digit level) are similar.

Table 3: Annual Sector Transitions and Failed Re-Accessions

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

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

are reallocated from traded-goods sectors with a comparative disadvantage to traded-goods sectors with comparative advantage so that the largest fraction of reallocated workers should move to the high-quintile sectors (columns 4 and 5) from every separation sector. Only in the aggregate of all separations (last row), including reallocations that failed at the annual horizon before, is there a higher absorption rate into comparative advantage sectors (especially column 4).

The dominant fraction of workers with displacement from a traded-goods plant, about a third, finds employment in nontraded-goods sectors (column 6). And almost as many workers with displacements from a traded-goods sector, roughly another third, are not rehired into any formal job within a year (column 7). Three quarters of the workers who are not reallocated at the annual horizon but who find reemployment in subsequent years move to the nontraded sector (second-to-last row) and, among the traded sectors, mostly into high-quintile sectors. At the national level, and in metropolitan areas on average, these statistics remain remarkably similar throughout the 1990s. In São Paulo state, however, reabsorption rates for displaced male worker in the traded-goods sector begin to increase in the second half of the 1990s, while reabsorption in the

Table 4: Annual Transitions Across Firms

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

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

nontraded sector drops.[7] These patterns are consistent with the idea that temporary work status changes out of formality (recorded as failures here), and jobs in the nontraded sector, provide a buffer in the prolonged reallocation process after trade reform.

Brazil's real exchange rate overvaluation during the 1990s arguably contributes to the expansion of the nontraded sector. Similarly, FDI inflows and the privatization of utilities and services firms in the second half of the 1990s as well as more frequent outsourcing of services jobs to suppliers (*terceirização*) might contribute to Brazil's nontraded sector expansion. (Regression analysis will control for FDI and privatization at the sector level, and for occupations susceptible to outsourcing at the job level.)

Trade models with heterogeneous firms predict that more productive firms become exporters more frequently, expand after a multilateral trade liberalization (such as the *Mercosur* agreement), and absorb workers (e.g. Melitz 2003). Table 4 reports transitions of displaced formal-sector workers to other formal-sector jobs at the annual horizon in 1990-91 and 1996-97, separately for nonexporting and exporting firms (*SECEX* exporter data are linked to *RAIS* at the firm level). Among the (one million) prime-age male workers who experience displacement from a formal job in 1990 and are rehired into a formal-sector job by December 1991, fewer than one in eleven workers move to a job at an exporting firm (columns 2 and 3). Accessions of displaced workers to exporters remain largely unaltered between 1990 and 1997 (columns 2 and 5) and their

[7] In-depth statistics are available online from *econ.ucsd.edu/muendler/research*.

magnitude is a small factor in the reabsorption of displaced workers. To summarize, a reason for Brazil's failing labor reallocation appears to be that both comparative advantage sectors and exporters fail to absorb displaced workers for a period of at least eight years after trade reform.

Theories. Reallocation frictions affect the predictions of classic trade models. Mussa (1978) introduces adjustment costs of factor employment into the Heckscher-Ohlin-Samuelson model and shows that factor owners' rational expectations set the economy on a path that maximizes the present value of output but that the long-run equilibrium critically depends on the adjustment technology.[8] Brecher (1974) considers factor-price distortions in the Heckscher-Ohlin-Samuelson model and provides an example of the eradication of gains from trade through incomplete labor reallocation because of inflexible real wages.

Introducing search frictions and unemployment into the Heckscher-Ohlin-Samuelson model, Davidson et al. (1999) show that the country with a more efficient search technology has a comparative advantage in the sector with high unemployment and vacancy rates. In a life-cycle model of goods, Saint Paul (1997) shows that countries with low factor-displacement costs exhibit a comparative advantage in industries with high aggregate demand volatility. Similarly for a Ricardian model with firm heterogeneity, Cunat and Melitz (2006) show that countries with flexible labor markets (where real wages adjust without negotiation) specialize in industries with much volatility in productivity.

It remains to establish our empirical claim that it is indeed the heightened trade exposure of Brazilian manufacturers that induces more worker displacements, hampers formal-sector labor reallocation, and increases transitions to informality. After describing our data (Section 3) and method (Section 4), we return to those issues in Section 5 with worker-level estimation.

[8]Temporary factor immobility in the Heckscher-Ohlin-Samuelson model causes the returns of one sector's mobile and immobile factor to initially move in the same direction, contrary to the Stolper-Samuelson result for the long term (Mayer 1974, Neary 1978). Our and prior empirical research shows, however, that the main concern is not the lock-in of labor: worker separations typically increase in the wake of trade reform. It is rather the paucity of reaccessions after separation that characterize the failed reallocation process.

3 Data

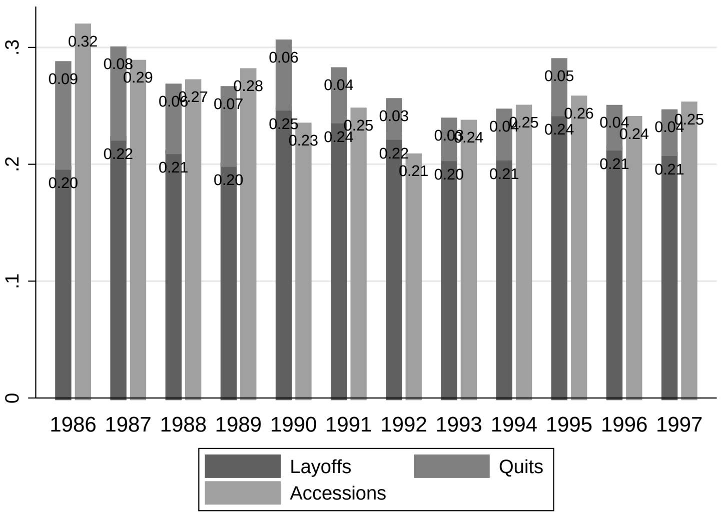
Of foremost concern for labor reallocation is the group of prime-age workers with no current formal-sector employment. It is the group of workers *to be reallocated.* Displacements from formal jobs add to the group, re-accessions into formal jobs shrink the group. We use mainly two data sources to circumscribe the group of workers to-be-reallocated. The metropolitan household survey *PME* provides direct information on workers with no formal-sector employment and, being a rotating panel, includes a single one-year span to observe transitions out of and into formality. The national labor force records *RAIS* include, by law, all formally employed workers, identify their plants, report the plant's sector, can be linked to firm information such as export status, naturally cover migrants,[9] and track the workers over time so that worker-fixed effects become estimable. By design, however, workers with no current formal-sector employment are not in *RAIS*.

To investigate with *RAIS* the size of the group of workers to-be-reallocated, we consider the group's two margins. We look at group additions: the separations from formal jobs, pushing workers into the group to-be-reallocated. And we watch group removals: the job accessions of prime-age male workers, or lifts out to formal employment. Figure 3 shows separation and accession rates for prime-age males with formal jobs in the manufacturing sector. Between 1986 and 1989, accessions mostly exceed separations (layoffs and quits) so that the group of prime-age workers to-be-reallocated does likely not grow because of labor turnover.[10] Starting in 1990, the converse prevails. Separations mostly exceed accessions so that the group to-be-reallocated is growing, resulting in longer durations of successful reallocations and a larger share of failed reallocations.

Worker data. Our linked employer-employee data derive from the labor force records *RAIS* (*Relação Anual de Informações Sociais* of the Brazilian labor ministry *MTE*). *RAIS* is a nationwide, comprehensive annual record of workers formally employed in any sector (including the public sector). The full data include 71.1 million workers (with 556.3 million job spells) at 5.52 million plants in 3.75

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

[10]Brazil's population growth may add to the group to-be-reallocated.



Source: *RAIS* 1986-97 (1% random sample), male workers nationwide, 25 to 64 years old, with employment in subsector *IBGE* manufacturing (highest paying job if many). Separation and accession rates exclude transfers, deaths, and retirements and are relative to totals of first and last observed employments in a given year.

Figure 3: **Separations and accessions in manufacturing**

million firms over the 16-year period 1986-2001. Every observation is uniquely identified by the worker ID (*PIS*), the plant ID (of which the firm ID is a systematic part), the month of accession, and the month of separation. Relevant worker information includes tenure at the plant, age, gender, and educational attainment; job information includes occupation and the monthly average wage; establishment information includes sector, municipality, and public-private ownership categories. We relegate further details on the data to the Appendix.

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

[11] We also draw a five-percent sample of all worker IDs that ever appear in a metropolitan area for direct comparisons to *PME*, and follow the workers nationwide. In addition, we repeat the calculation of statistics and estimation with the *CNAE* sector classification (roughly comparable to the *NAICS* four-digit level), which is available since 1995, by using

We keep only male workers, age 25 to 64, in order to be little affected by labor supply changes and to capture mostly workers past their first entry into the labor force. For most statistics, we remove multiple jobs and retain a worker's highest paying job observation. When we infer separations and accessions, we exclude transfers across plants (at different locations) within the same firms, as well as retirements and deaths on the job. An accession is defined as a worker's hiring into the first employment in the calendar year; reference observations for accession rates and estimation are employments with no reported accession in the year. Conversely, a separation is a worker's quit or layoff from the last employment in the calendar year; reference observations for separation rates and estimation are employments with no reported separation in the year. Quits are infrequent compared to layoffs (Figure 3) and not clearly distinguishable in practice. So we mostly consider separations as a single category. We construct plant-level labor force information by aggregating *RAIS* to the plant level.

Linked plant and firm data. We use annual customs office records from *SECEX* (*Secretaria de Comércio Exterior*) on exports for 1990 through 1998. We set the indicator variable for a firm's exporting status to one when *SECEX* records show exports of any product from the firm in a given year.[12] We link the export-status indicator to *RAIS* at the firm level.

To link firm-level total factor productivity (TFP) information to *RAIS* for a robustness check, we use the annual manufacturing firm survey *PIA* (*Pesquisa Industrial Anual* from Brazil's census bureau *IBGE*) for 1986-98. *PIA* is a representative sample of all but the smallest manufacturing firms. We obtain log TFP measures from Olley and Pakes (1996) estimation at the *Nível 50* sector level (Muendler 2004). *IBGE*'s publication rules allow data from *PIA* to be withdrawn in the form of tabulations of cells with at least three firms. We construct firm mini-groups using three-firm random combinations drawn from within each *Nível 50* sector, headquarter location, and possible sequence of consecutive calendar years. We assign a *PIA* firm to one and only one mini-group observation. A single four- or five-firm mini-group is defined within a sector-year-location cell when the number of firms in the sector-year-location cell is not divisible by three. For each three-to-five-firm mini-group observation, we

a sample of surviving plants through 1995. The figures, tables and estimates are online at *econ.ucsd.edu/muendler/research.*

[12] We do not use sales thresholds to define the export indicator because sales information is only available for a random subsample of (*PIA*) firms. Our regressions control for establishment employment, so exports per worker would not add information.

calculate mean log TFP but retain the firm identifiers behind the mini group—permitting the linking to *RAIS*.

Complementary household survey data. The Brazilian monthly employment survey *PME* (*Pesquisa Mensal de Emprego*, from *IBGE*) cannot be linked to establishments or firms but provides details on work status. *PME* data derive from a random sample of households in six metropolitan areas (São Paulo, Rio de Janeiro, Belo Horizonte, Porto Alegre, Salvador, Recife). The data are collected from a rotating panel similar to *PSID* in the U.S. *PME* follows households for 16 months, with an eight-month interval after the fourth interview.[13] Changes to the sample design adversely affect worker panels starting in odd years. So, we use only individuals whose first survey occurs in 1986, 1988, 1990, 1992, 1994, 1996 or 1998. As with *RAIS*, we restrict our sample to prime-age male workers. We only trace changes in the work status between the fourth and the 8th interviews for each household member so that we can control for the individual's work status during the three months prior to the fourth interview.

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

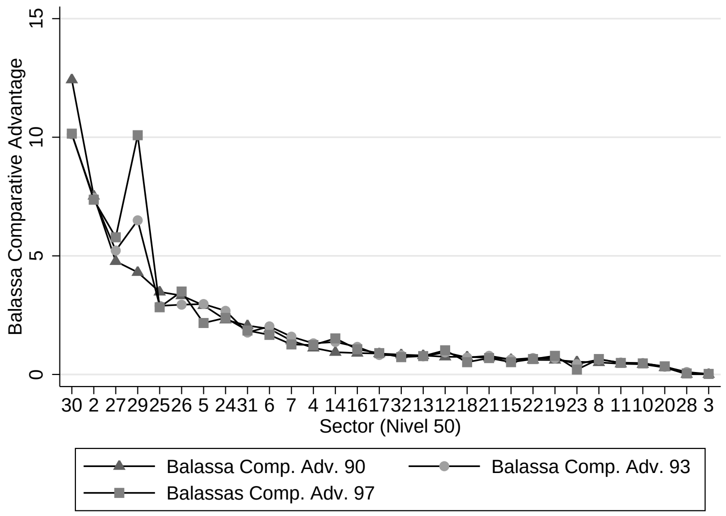
Sector data. We construct sector-level variables from various sources. We use UN Comtrade trade data for 1986-98 to construct Balassa (1965) comparative advantage measures for Brazil. Balassa (1965) comparative advantage of sector i in year t is

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

where $X_{i,t}$ are exports.

The Balassa (1965) statistic is an especially adequate measure of comparative advantage if sector rankings remain stable over time. Indeed, regressions of $BADV_{i,t}$ on product and input tariffs, year and sector indicators show that

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



Sources: UN Comtrade 1986-98. Sectors at *Nível 50* ranked by Balassa comparative advantage FE (for sector definitions see Table 17, p. 42). Estimates of Balassa comparative advantage fixed effects (FE) from sector-fixed effects regression on output tariffs, input tariffs and year indicators.

Figure 4: **Balassa Comparative Advantage**

$BADV_{i,t}$ is highly sector specific, statistically unrelated to tariffs conditional on sector effects, and time-invariant. Year indicators are neither individually nor jointly different from zero at common significance levels.[14] Figure 4 ranks manufacturing sectors by their estimated sector coefficients from a linear sector fixed-effects regression on product and input tariffs and year indicators. The Figure illustrates that Brazil's sectors of revealed comparative advantage remain largely the same throughout 1990-97. With the exception of sector 27 (meat processing), which advances to stronger comparative advantage, comparative advantage changes hardly at all. This time invariance is also consistent with the lacking permeability of sectors to cross-sector labor reallocations (Table 3): sectors with different degrees of comparative advantage are distinct.

Our main instrumental variables for firm-level export status are imports into Brazil's export destinations from countries other than Brazil, weighted with Brazil's sectoral export volumes in the base year 1990. We use WTF (*NBER*)

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

data on bilateral trade 1990-98 to construct the instruments by subsector *IBGE* and seven world regions: Asia-Pacific Developing countries (APD), Central and Eastern European countries (CEE), Latin American and Caribbean countries (LAC), North American countries (NAM excluding Mexico), Other Developing countries (ODV), Other Industrialized countries (OIN), and Western European countries (WEU). From *SECEX* exports data we calculate sector-specific weights for each foreign destination country in 1990 (using *RAIS* sector information for the *SECEX* exporters). We then calculate aggregate imports into each foreign country, excepting imports from Brazil, and weight the country aggregates with Brazilian export volumes by sector and destination. We finally aggregate the sector-weighted country totals to seven world regions and obtain seven foreign import-demand instruments that vary by sector and year.

Additional instruments, for tariffs and import penetration rates, are real exchange rate components at the sector-year level: the USD exchange rate and sectoral price levels in the U.S. and the EU. We relegate details on these instruments and additional sector variables to the Appendix.

Descriptive statistics. We focus our regression analysis on the manufacturing sector. Table 5 summarizes the linked employer-employee data for the separation sample. The accession sample has similar statistics. Annual separation rates (from the separation sample) and accession rates (from the accession sample) are close to thirty percent (column 1). This is almost double the magnitude of industrialized-country labor markets. Unconditionally, mean separation and mean accession rates are larger in comparative-advantage sectors in the fifth quintile than in the least-advantage sectors (columns 3 and 4).

Roughly half the manufacturing labor force is employed at a firm that exports at least one product in a given year (column 1). As we expect for a country with a history of import substitution, the comparative-advantage sectors in the first quintile exhibit higher tariffs (but also higher import penetration rates) than Brazil's sectors with a comparative advantage (columns 3 and 4). Interestingly, a larger portion of the labor force is employed at exporters in sectors with the least advantage than at exporters in the strongest-advantage sectors. At least two explanations are consistent with this finding. First, exporters in the least advantage sector exhibit lower labor productivity and larger employment. Second, there is a larger number of exporters in the least-advantage sectors with small-scale exports than in the strongest-advantage sectors. Both explanations are borne out in our data. We take firm-level theory as a strict guide, however,

Table 5: *RAIS* Summary Statistics for Manufacturing

	All sectors		Comp. adv. quintile 1st	Comp. adv. quintile 5th
	Mean	Std.Dev.	Mean	Mean
	(1)	(2)	(3)	(4)
Outcomes				
Indic.: Separation	.282	.450	.272	.314
Indic.: Layoff	.245	.430	.242	.262
Indic.: Quit	.026	.160	.023	.031
Indic.: Accession	.292	.455	.264	.326
Main covariates				
Balassa (1965) Comp. Adv.	1.450	1.047	.537	3.223
Exporter Status	.495	.500	.556	.438
Product Market Tariff	.193	.103	.232	.174
Intm. Input Tariff	.146	.077	.190	.105
Import Penetration	.064	.052	.100	.031
Sector-level covariates				
FDI Flow (USD billion)	.110	.334	.076	.263
Herfindahl Index (sales)	.089	.056	.066	.083
Share: Jobs at private firms	.955	.019	.949	.966
Plant-level covariates				
Log Employment	5.148	1.952	5.110	5.551
Share: Middle School or less	.745	.219	.684	.815
Share: Some High School	.182	.159	.223	.137
Share: White-collar occup.	.264	.211	.306	.241
Worker-level covariates				
Tenure at plant (in years)	.952	1.208	.992	.778
Pot. labor force experience	25.276	9.971	24.373	26.116
Middle School or less	.785	.411	.727	.854
Some High School	.151	.358	.187	.108
Some College	.020	.141	.029	.012
College Degree	.038	.191	.053	.021
Prof. or Manag'l. Occ.	.085	.278	.100	.069
Tech'l. or Superv. Occ.	.082	.274	.108	.061
Unskilled Wh. Collar Occ.	.070	.255	.078	.079
Skilled Bl. Collar Occ.	.636	.481	.614	.646
Unskilled Bl. Collar Occ.	.102	.303	.079	.120
Indic.: Outsourceable job	.252	.434	.253	.234

Source: *RAIS* 1990-98 (1% random estimation sample), male workers nationwide, 25 to 64 years old, with manufacturing job. Statistics based on separation sample, except for accession indicator (146,800 observations in separation, 112,971 in accession sample). Sector information at subsector *IBGE* level.

and do not introduce degrees of exporting status.

Plants in the strongest-advantage sectors attract more FDI, are more concentrated, are more frequently privately owned, and larger in employment than in the least-advantage sectors. Workers have shorter tenure in the strongest-advantage sectors, consistent with higher labor turnover in the strongest-advantage sectors, but on average almost two years more labor force experience. As the Heckscher-Ohlin-Samuelson model would predict, Brazil, a relatively low-skill abundant country, employs a larger share of less educated workers in the strongest-advantage sectors than in the least-advantage sectors, and offers less skill-intensive occupations in the strongest-advantage sectors. Occupations in the strongest-advantage sectors are less susceptible to outsourcing than in the least-advantage sectors.

4 Methods

Wage-taking plants adjust employment through worker separations and accessions. Consider the probability that an employer-employee match is terminated or formed, conditional on a worker-fixed effect that is observable to the employer and known to the employee:

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

where the indicator $\sigma_{i,t}$ takes a value of unity for the outcome (accession or separation) for worker i at time t and zero otherwise. α_i is a worker-fixed effect; $\mathbf{z}_{S(J(i)),t}$ is a vector of sector-level covariates of the worker's displacing or hiring sector $S(J(i))$; $\mathbf{y}_{J(i),t}$ is a vector of plant-level covariates of worker i's displacing or hiring plant $J(i)$; $\mathbf{x}_{it}$ is a vector of covariates that are worker or job specific, or both; and β_z, β_y, β_x are coefficient vectors. There is an unobserved residual to the termination and formation of employer-employee matches. Under suitable assumptions on the error term, probability model (1) becomes a conditional logit model for worker panels (FE cLogit) which we fit, as is common, using conditional maximum likelihood estimation because α_i and β are inconsistent under full maximum likelihood estimation (McFadden 1974).

The FE cLogit estimator operates on workers who experience at least one separation or accession. This reduces sample size. However, the magnitude of labor turnover in Brazil, the overall sample size, and the sample time span of nine years ensure identification. Worker- and occupation-level covariates are identified from time variation within or across jobs under the FE cLogit estimator. Coefficients on occupation-level covariates are identified through workers who also switch occupation. Educational attainment changes little among prime-age males. We consequently drop education categories from the worker characteristics vector but keep educational workforce composition shares among the plant-level regressors. To assess the importance of worker-fixed effects in the termination and formation of employer-employee matches, we compare results to a conventional logit model with worker-clustered standard errors.

A Rivers and Vuong (1988) test for the significance of residuals from instrumental-variable prediction shows that that simultaneity between several regressors and the outcome is an empirical issue in our FE cLogit estimation. For we are mostly concerned with signs of coefficients and their statistical significance, and less interested in magnitudes of odds ratios, we repeat estimation with a linear worker-fixed effects regression (linear FE) under instrumental variables, analogous to (1), to remove simultaneity bias.

We use a multinomial logit (MNL) model to estimate transitions between work status categories from the set $\mathbb{S}$ of alternatives. Under MNL assumptions, an individual household member's probability to move to work status $\sigma_{i,t+1}$, conditional on present work status $\sigma_{i,t} = \sigma$, is

$$Pr(\sigma_{i,t+1}|\sigma_{i,t} = \sigma; \mathbf{x}, \mathbf{z}) = \frac{\exp\{\mathbf{z}_{S(i),t}\beta_z^\sigma + \mathbf{x}_{i,t}\beta_x^\sigma\}}{\sum_{\varsigma\in\mathbb{S}} \exp\{\mathbf{z}_{S(i),t}\beta_z^\varsigma + \mathbf{x}_{i,t}\beta_x^\varsigma\}}, \tag{2}$$

where $\mathbf{z}_{S(i),t}$ is a vector of sector-level covariates of the worker's displacing sector $S(i)$; $\mathbf{x}_{it}$ is a vector of covariates that are worker specific; and β_x^ς and β_z^ς are coefficient vectors for the future work status $\varsigma \in \mathbb{S}$. Coefficients are identified relative to a baseline work status at $t+1$. We use as the baseline category a worker's continuation in the present work status, $\sigma_{i,t+1} = \sigma_{i,t} = \sigma$. We obtain robust standard errors for all coefficients in the FE cLogit regressions, linear FE regressions (with instruments), and MNL regressions.

5 Estimates

We start with a direct look at the group of workers to-be-reallocated using the household survey *PME*. We then proceed to richer sets of predictors using *RAIS* linked employer-employee data by investigating formal-job separations and accessions: the margins that augment or shrink the group of workers to-be-reallocated.

5.1 Work status transitions

We estimate the MNL model for two types of household members: household members with a formal manufacturing job in the base year, and household members with an informal manufacturing job in the base year. Consider the workers with a formal manufacturing job first. The set of work status categories for our MNL estimator includes five alternatives: (1) the worker retains the formal manufacturing job or switches to a new formal job (not necessarily in manufacturing); (2) the worker moves to an informal job (not necessarily in manufacturing); (3) the worker moves to self-employment (not necessarily in manufacturing); (4) the worker moves to unemployment; and (5) the worker leaves the labor force. For workers with informal manufacturing jobs in the base year, alternative (1) becomes that the worker retains or moves to an informal job, and (2) that the worker transitions to a formal job (not necessarily in manufacturing).

We pool the *PME* household member observations at the annual horizon (between the fourth and eighth interview) over the sample period 1986-99, and control for year effects, location effects and worker gender. Aware of the importance of worker-level heterogeneity, we create an indicator variable that takes a value of unity if formality (informality) status lasted for less than three months prior to the fourth interview.

Table 6 presents predictions of transitions from formal manufacturing work status, controlling for year and location effects. Protective product market tariffs are associated with reduced probabilities of transitions from formal manufacturing employment to an alternative work status (except withdrawals from the labor force). Most striking, and contrary to findings by Goldberg and Pavcnik (2003) for sectoral data on Brazil, the odds of transitions from formal manufacturing employment to the informal sector (column 1), and to self-employment (column 2), are significantly higher in the presence of reduced product market tariff barriers. Reduced input tariffs predict significantly more transitions to unemployment. Competition-aggravating high input tariffs correlate positively with transitions into self employment, however, leaving the net effect of trade reform on self employment transitions ambiguous (column 2). This offsetting effect is not observed for informality transitions: reduced barriers to foreign competition predict significantly higher odds of informality, contrary to previous evidence.

Workers with stable formal-sector employment over the four months preceding the fourth interview are less likely to lose formality status, conditional a worker's potential labor force experience (not reported). Higher educational attainment mostly predicts less transitions out of formality, with the notable

Table 6: Work Status Transitions from Formal Employment

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

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

exception of college-educated workers who suffer relatively more displacements to informality.

Table 7 shows estimates of the reverse transition out of informal manufacturing employment and into different types of work status, among them formality. We control for year and location effects, and a worker's potential labor force experience. Elevated product tariffs significantly raise the odds that a worker transitions out of informality into formal employment, or any other category. For this margin, competition-aggravating high input tariffs have no statistically detectable effect. Workers with lasting informal-sector employment over the past four months are less likely to make it out of informality status. Compared to primary-school educated and illiterate workers, higher educational attainments significantly reduce the chances of a transition out of informality. Similar to the results for transitions out of formality, higher education levels seem to contribute to employment stability irrespective of work status.

Controlling for worker characteristics, trade reform has a statistically sig-

Table 7: WORK STATUS TRANSITIONS FROM INFORMAL EMPLOYMENT

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

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

nificant impact both on transitions from formality into informality and from informality into formality. The estimates are consistent with the hypothesis that workers suffer longer spells of informality when their sector is exposed to more severe foreign competition. We now turn to an analysis of worker, job, employer and sector covariates that predict separations from and accessions into formal-sector manufacturing jobs.

5.2 Formal-sector turnover

Table 8 presents worker-FE cLogit estimates of displacements from formal manufacturing jobs. Displacements are significantly more frequent in sectors with a stronger comparative advantage and at exporter firms. Heightened import penetration after trade reform predicts significantly higher displacement odds. The effect of tariffs on displacements is expectedly mixed, however. Pro-competitive product tariff cuts predict significantly higher displacement rates from formal

Table 8: Conditional Logit Estimates of Separations

	(1)	(2)	(3)	(4)	(5)
Balassa (1965) Comp. Adv.	.079				.168
	(.021)***				(.024)***
Exporter Status		.289			.283
		(.028)***			(.028)***
Product Market Tariff			-.111		-.710
			(.416)		(.426)*
Intm. Input Tariff			1.617		2.893
			(.633)**		(.678)***
Import Penetration				.770	1.247
				(.353)**	(.388)***
Sector-level covariates					
FDI Flow (USD billion)	-.025	-.012	-.017	-.013	-.048
	(.020)	(.020)	(.020)	(.020)	(.020)**
Herfindahl Index (sales)	-.371	-.516	-.397	-.655	-.350
	(.317)	(.316)	(.329)	(.325)**	(.343)
Plant-level covariates					
Log Employment	-.343	-.370	-.341	-.339	-.377
	(.011)***	(.011)***	(.011)***	(.011)***	(.011)***
Share: Middle School or less	-.746	-.654	-.715	-.713	-.659
	(.131)***	(.131)***	(.131)***	(.131)***	(.132)***
Share: Some High School	-.439	-.387	-.435	-.438	-.388
	(.148)***	(.148)***	(.147)***	(.147)***	(.148)***
Share: White-collar occup.	.725	.704	.743	.742	.695
	(.075)***	(.074)***	(.075)***	(.074)***	(.075)***
Worker-level covariates					
Tenure at plant (in years)	1.367	1.350	1.362	1.363	1.351
	(.036)***	(.036)***	(.036)***	(.036)***	(.036)***
Pot. labor force experience	.006	.006	.006	.006	.006
	(.002)**	(.002)**	(.002)**	(.002)**	(.002)**
Tech'l. or Superv. Occ.	-.070	-.067	-.076	-.071	-.077
	(.068)	(.068)	(.068)	(.068)	(.068)
Unskilled Wh. Collar Occ.	-.257	-.252	-.260	-.256	-.263
	(.067)***	(.067)***	(.067)***	(.067)***	(.067)***
Obs.	145,418	145,418	145,418	145,418	145,418
Pseudo R^2	.148	.149	.148	.148	.150

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

Table 9: Conditional Logit Estimates of Accessions

	(1)	(2)	(3)	(4)	(5)
Balassa (1965) Comp. Adv.	.041				-.016
	(.017)**				(.020)
Exporter Status		-.449			-.439
		(.027)***			(.027)***
Product Market Tariff			1.309		1.248
			(.379)***		(.393)***
Intm. Input Tariff			-3.265		-3.078
			(.540)***		(.598)***
Import Penetration				-.519	.203
				(.320)	(.355)
Sector-level covariates					
FDI Flow (USD billion)	.039	.046	.056	.047	.058
	(.022)*	(.021)**	(.021)***	(.021)**	(.022)***
Herfindahl Index (sales)	-.346	-.343	-.795	-.275	-.788
	(.268)	(.268)	(.282)***	(.277)	(.297)***
Plant-level covariates					
Log Employment	-.191	-.140	-.190	-.189	-.141
	(.008)***	(.009)***	(.008)***	(.008)***	(.009)***
Share: Middle School or less	.949	.858	.942	.950	.852
	(.107)***	(.105)***	(.107)***	(.107)***	(.105)***
Share: Some High School	.742	.669	.741	.742	.670
	(.124)***	(.122)***	(.124)***	(.124)***	(.122)***
Share: White-collar occup.	-.676	-.615	-.680	-.673	-.622
	(.067)***	(.067)***	(.067)***	(.067)***	(.067)***
Worker-level covariates					
Prof. or Manag'l. Occ.	-.799	-.805	-.799	-.798	-.806
	(.068)***	(.068)***	(.068)***	(.068)***	(.068)***
Tech'l. or Superv. Occ.	-.602	-.609	-.595	-.601	-.603
	(.064)***	(.064)***	(.064)***	(.064)***	(.064)***
Unskilled Wh. Collar Occ.	-.491	-.498	-.489	-.490	-.497
	(.061)***	(.062)***	(.062)***	(.061)***	(.062)***
Skilled Bl. Collar Occ.	-.418	-.414	-.414	-.417	-.411
	(.032)***	(.032)***	(.032)***	(.032)***	(.032)***
Obs.	112,971	112,971	112,971	112,971	112,971
Pseudo R^2	.036	.040	.037	.036	.041

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

jobs, whereas the reductions of competition-aggravating input tariff barriers alleviate competition and significantly reduce displacement rates.[15] FDI inflows stabilize employment by reducing displacement rates.

Before discussing plant and worker-level variables, we turn to predictions for the converse margin: accessions to formal manufacturing jobs. Table 8 presents the worker-FE cLogit estimates. Mirroring the sign from the separation regression, accessions are less frequent in sectors with a stronger comparative advantage once we condition on other trade-related variables (column 5). The coefficient is not statistically significant at conventional levels in this regression but becomes significant when controlling for higher-order interactions between trade variables (Table 15). Exporters exhibit significantly lower accession rates, mirroring their higher displacement rates. The effect of tariffs on accessions is mixed again. Pro-competitive product tariff cuts predict significantly fewer accessions to formal jobs, whereas the reductions of competition-aggravating input tariff barriers alleviate competition and significantly raise accession rates. Heightened import penetration after trade reform predicts is not associated with significantly different accession odds. Mirroring separations, FDI inflows stabilize employment by raising accession rates. More concentrated industries exhibit significantly fewer accessions.

To summarize, increased import penetration of the Brazilian manufacturing sector triggers significantly more displacements. But neither comparative-advantage sectors nor exporters provide the arguably expected absorption for reallocation. In the opposite, comparative-advantage sectors displace workers significantly more frequently than other sectors and tend to hire fewer workers. Similarly, exporters displace workers significantly more frequently and absorb workers significantly less frequently than non-exporters.

Larger plants exhibit less turnover: they displace fewer (Table 8) and they hire fewer workers (Table 9). At the displacement margin, plants with more highly educated workforces (beyond middle and high school attainment) and plants with more white-collar occupations shed less workers, while the mirror picture arises for accessions. Plants with more highly educated workforces (beyond middle and high school attainment) and plants with more white-collar occupations hire more frequently. Among the least frequently displaced white-collar

[15] We favor the interpretation of elevated input tariffs as anti-competitive measures. If their removal were mainly a proxy for access to foreign intermediates and capital goods, displacement coefficients would have to show the opposite sign unless those goods are complements to Brazilian labor. Firms with labor productivity advances, however, reduce employment (Table 1), rendering foreign-factor and technology complementarities unlikely.

workers are those in unskilled white-collar occupations, whereas unskilled blue-collar workers experience the relatively most frequent accessions. A worker's tenure at the plant and labor market experience predict a significantly higher displacement probability. Year effects (not reported) are significant at the one-percent level and show both a strictly monotonic increase in displacements and a strictly monotonic drop in accessions.

6 Robustness checks

Our main hypothesis is that heightened trade exposure triggers displacements and that neither comparative-advantage sectors nor exporters absorb the displaced workers. We now assess the plausibility of alternative hypotheses.

Simultaneity bias. Firms decide exporting status and labor turnover simultaneously. In a similar vein, Brazil's trade liberalization programme intentionally targets low-efficiency sectors with the strongest tariff reductions and foreign competitors more likely enter Brazil's low-efficiency sectors. But initially protected sectors with a comparative disadvantage exhibit less labor turnover (Table 5). So, the potential simultaneity of exports with firm-level labor turnover and the simultaneity of tariffs and import-penetration rates with sectoral labor turnover are a concern. We use destination-country imports from other source-countries than Brazil, weighted with sector-level export volumes from Brazil in the base year, as well as components of the sectoral real exchange rate as instruments.[16] Predicting export status, tariffs and import penetration with these instruments at the sector-year level, and including both predicted values and residuals in FE cLogit estimation, shows that coefficients on the residuals are highly significant and that simultaneity is an empirical issue (Rivers and Vuong 1988). To assess the magnitude of the potential simultaneity bias, and check for potential sign reversals, we resort to linear FE regressions of separation and accession indicators on the same set of predictors as before.

We predict potentially simultaneous variables with instruments. Table 10 shows results of these first stage regressions (we do not report results from input-tariff regressions, which are similar to estimates in columns 2 and 5). The regressions are weighted by employment observations in the separation and accession

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

Table 10: First-stage Predictions

	Separations			Accessions		
	Exp. Status	Prd. Mkt. Tariff	Imp. Pen.	Exp. Status	Prd. Mkt. Tariff	Imp. Pen.
	(1)	(2)	(3)	(4)	(5)	(6)
Instruments						
World imports APD	3.530	-2.279	-.004	3.847	-2.121	.386
	(.789)***	(.097)***	(.053)	(.975)***	(.111)***	(.065)***
World imports CEE	43.555	-33.877	-16.621	38.918	-26.910	-17.062
	(4.341)***	(.534)***	(.293)***	(5.551)***	(.635)***	(.370)***
World imports LAC	-4.784	14.267	4.756	-2.023	14.042	4.863
	(1.035)***	(.127)***	(.070)***	(1.319)	(.151)***	(.088)***
World imports NAM	-2.363	-.651	-1.670	-2.482	.376	-1.991
	(.525)***	(.065)***	(.035)***	(.662)***	(.076)***	(.044)***
World imports ODV	-2.127	-5.736	.312	-1.391	-5.273	-.140
	(.763)***	(.094)***	(.052)***	(.977)	(.112)***	(.065)**
World imports OIN	4.175	-9.099	-5.676	3.989	-10.353	-5.340
	(.957)***	(.118)***	(.065)***	(1.181)***	(.135)***	(.079)***
World imports WEU	13.971	2.157	1.949	14.443	1.468	2.095
	(.461)***	(.057)***	(.031)***	(.564)***	(.065)***	(.038)***
PPI Idx. EU	.706	-.929	.112	.975	-.940	.052
	(.115)***	(.014)***	(.008)***	(.144)***	(.016)***	(.010)***
PPI Idx. NAM	.411	.850	-.120	.475	.802	-.200
	(.106)***	(.013)***	(.007)***	(.138)***	(.016)***	(.009)***
USD Exch. Rate	.106	-.211	.011	.081	-.252	-.014
	(.025)***	(.003)***	(.002)***	(.032)**	(.004)***	(.002)***
Exogenous covariates						
Balassa Comp. Adv.	-.020	-.026	-.022	-.024	-.027	-.022
	(.003)***	(.0003)***	(.0002)***	(.003)***	(.0004)***	(.0002)***
FDI Flow	.002	.014	.004	.0001	.014	.005
	(.003)	(.0004)***	(.0002)***	(.004)	(.0004)***	(.0003)***
Herfindahl Index	.329	.048	.053	.253	-.026	.098
	(.044)***	(.005)***	(.003)***	(.054)***	(.006)***	(.004)***
Log Employment	.052	.003	-.0009	.050	.003	-.0007
	(.002)***	(.0002)***	(.0001)***	(.002)***	(.0002)***	(.0001)***
Middle Sch. Share	-.172	.008	-.007	-.184	.007	-.009
	(.016)***	(.002)***	(.001)***	(.017)***	(.002)***	(.001)***
Some High Sch. Share	-.062	-.002	.003	-.092	-.005	.002
	(.019)***	(.002)	(.001)*	(.021)***	(.002)**	(.001)
White-collar Share	.060	.006	-.002	.056	.004	-.002
	(.010)***	(.001)***	(.0007)**	(.012)***	(.001)***	(.0008)**
R^2	.104	.856	.612	.121	.849	.619
F statistic (instr.)	13.529	14347.77	475.52	23.768	12720.89	310.409

Sources: *WTF* (*NBER*) bilateral import data 1990-98; sector data 1990-98 from various sources at subsector *IBGE* level; *RAIS* 1990-98 labor force information; *SECEX* exporter information 1990-98. Weighted regressions using worker-sample observations (as in Table 8 for separations, Table 9 for accessions), controlling for year effects. Annual sector-weighted world imports, coefficients rescaled to imports in USD trillion. Robust standard errors in parentheses: * significance at ten, ** five, *** one percent.

Table 11: Fixed-Effects and Instrumental-Variable Estimates

	Separations			Accessions		
		OLS-FE			OLS-FE	
	Cdl. logit		IV	Cdl. logit		IV
	(1)	(2)	(3)	(4)	(5)	(6)
Balassa Comp. Adv.	.168	.034	.035	-.016	-.003	-.007
	(.024)***	(.004)***	(.005)***	(.020)	(.004)	(.005)
Exporter Status	.283	.060	.011	-.439	-.091	-.383
	(.028)***	(.005)***	(.123)	(.027)***	(.005)***	(.152)**
Product Market Tariff	-.710	-.160	.013	1.248	.285	.219
	(.426)*	(.065)**	(.139)	(.393)***	(.078)***	(.160)
Intm. Input Tariff	2.893	.599	.552	-3.078	-.687	-.855
	(.678)***	(.099)***	(.241)**	(.598)***	(.118)***	(.295)***
Import Penetration	1.247	.167	.186	.203	.114	.462
	(.388)***	(.060)***	(.138)	(.355)	(.068)*	(.162)***
Obs.	145,418	145,418	145,418	112,971	112,971	112,971

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

samples. Almost every instrumental variable is a statistically significant predictor of the potentially simultaneous variables at the one-percent level. There is no evidence of weak instruments: F statistics from joint significance tests on the instruments vary between 14 and 14,000. We highlight a few coefficient estimates. Higher producer prices in the U.S. and Europe, as well as a weaker Brazilian Real, predict significantly more frequent exporting status. Exporters have larger and more educated workforces, and offer more white-collar jobs. As discussed in the data Section 3 earlier, employment-weighted exporting frequencies are higher in sectors with a weaker comparative advantage because exporters in the least advantage sector exhibit lower labor productivity and larger employment and because there is a larger number of exporters with small-scale exports in the least-advantage sectors. Product tariffs are lower in comparative-advantage sectors.

Table 11 compares the FE cLogit estimates on separations and accessions (from Tables 8 and 9) to those from linear FE regressions without and with

Table 12: Firm-level Labor Productivity

	Separations			Accessions		
	Full smpl.	*PIA* smpl.		Full smpl.	*PIA* smpl.	
	(1)	(2)	(3)	(4)	(5)	(6)
Balassa Comp. Adv.	.168	-.006	-.009	-.016	-.017	-.012
	(.024)***	(.074)	(.075)	(.020)	(.060)	(.060)
Exporter Status	.283	.030	.030	-.439	-.292	-.294
	(.028)***	(.076)	(.076)	(.027)***	(.075)***	(.075)***
Prod. Market Tariff	-.710	1.267	1.255	1.248	-.303	-.256
	(.426)*	(.987)	(.988)	(.393)***	(.955)	(.959)
Intm. Input Tariff	2.893	.381	.438	-3.078	-1.303	-1.513
	(.678)***	(1.608)	(1.610)	(.598)***	(1.375)	(1.376)
Import Penetration	1.247	1.322	1.278	.203	.497	.621
	(.388)***	(1.000)	(1.002)	(.355)	(1.098)	(1.099)
Log Labor Prod.			.008			-.020
			(.008)			(.007)***
Obs.	145,418	40,335	40,335	112,971	20,183	20,183
Pseudo R^2	.150	.335	.335	.041	.089	.090

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

instrumental-variable predictions. Both signs and significance coincide between FE cLogit (columns 2 and 4) and linear FE regressions (columns 3 and 5). When instrumenting the simultaneous variables—export status, tariffs, and import penetration—, there is not one significant sign change (comparing columns 2 and 3, and columns 5 and 6). A single coefficient estimate, on product tariffs in displacements, shows an insignificant sign change. Our main hypothesis, that heightened trade exposure triggers displacements and that neither comparative-advantage sectors nor exporters absorb displaced workers, is not rejected. Import penetration is positively associated with displacements, exporters tend to displace more and hire significantly fewer workers.

Firm-level labor productivity. Theory predicts and *PIA* data show for Brazil that exporters are more productive than non-exporters. We use the subsample of *PIA* firms, linked to *RAIS* to assess whether exporting status predicts more separations and less accessions beyond the productivity difference to non-exporters. Table 11 compares the FE cLogit estimates on separations and accessions (from Tables 8 and 9) to FE cLogit on the linked *PIARAIS* subsample without and with firm-level log labor productivity as a covariate. Trade-variable are not significant predictors of displacements in the random subsample of medium-sized to large manufacturers but point estimates with the log labor productivity control are close to those without log labor productivity. At the accession margin, both exporting status and log labor productivity are separately associated with significantly lower hiring rates. This confirms our empirical claim that exporters tend to displace more and hire significantly fewer workers.

Concomitant reforms and worker heterogeneity. Though the labor market reforms of 1988 precede trade reform, they might affect sectors with differently composed workforces to varying degrees and interact with trade reform in a way that erroneously attributes labor turnover to the trade regime. Similarly, the privatization of state-owned businesses and the progressing outsourcing of jobs to specialized service suppliers can differentially affect sectors and lead us to erroneously attribute labor turnover to the trade regime. We turn to the plausibility of these competing explanations.

For trade variables to exhibit a spurious relation with the labor-turnover outcomes of concomitant reforms, the sectoral workforce composition would have to matter for the impact of trade on labor turnover. We investigate this implication by repeating the FE cLogit for workers with varying educational attainment. Table 11 contrasts the FE cLogit estimates on separations and accessions (from Tables 8 and 9) to FE cLogit estimates for three levels of education: workers with no or some primary school education (column 2), workers with some or complete high school education (column 3), and workers with some or complete college education (column 4). Though magnitudes vary to a degree, the sign and significance patterns are strikingly similar across all three education groups. Tariffs have a somewhat stronger impact on workers with more than primary schooling. Point estimates most closely related to our main claims—that heightened import penetration triggers displacements, and that comparative-advantage sectors and exporters tend to displace more and hire fewer workers—, are close across workers of different education levels. So, concomitant reforms do not likely interact with effects of trade reform on labor turnover.

We take reforms one by one to further probe the plausibility of erroneous at-

Table 13: COMPLEMENTARY CONDITIONAL LOGIT ESTIMATES

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

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

tributions to trade reform. Labor market reform precedes trade reform, so sector fixed effects should account for its potential differential impact. Including sector effects at the subsector *IBGE* level in the regression (column 5) expectedly turns the coefficient on comparative advantage, which exhibits hardly any time variation, insignificant. For the other trade regressors, however, coefficient estimates increase in absolute value. This renders erroneous attribution of labor-market regulations to trade reform little plausible.

Privatization is partly observed in *RAIS*. Plants report their ownership type since 1995. We calculate the share of the workforce employed at privately-owned plants in the sector workforce and assign the 1995 values to prior years. The

Table 14: LOGIT ESTIMATES OF SEPARATIONS, 1986-98

	Separations		Accessions	
	Cdl. logit	Logit	Cdl. logit	Logit
	(1)	(2)	(3)	(4)
Balassa (1965) Comp. Adv.	.168	.125	-.016	.184
	(.024)***	(.007)***	(.020)	(.007)***
Exporter Status	.283	-.017	-.439	-.508
	(.028)***	(.012)	(.027)***	(.013)***
Product Market Tariff	-.710	-.491	1.248	-1.398
	(.426)*	(.170)***	(.393)***	(.198)***
Intm. Input Tariff	2.893	1.372	-3.078	2.226
	(.678)***	(.255)***	(.598)***	(.301)***
Import Penetration	1.247	-.330	.203	-1.001
	(.388)***	(.140)**	(.355)	(.165)***
Obs.	145,418	293,369	112,971	293,137
Pseudo R^2	.150	.050	.041	.078

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

Cardoso government pursues privatization forcefully since 1995 so that the 1995 privatization share is a useful approximation to the share in preceding years. Inclusion of the regressor hardly alters coefficient estimates for the trade variables (column 6). Finally, we construct a job-level indicator for an occupation's susceptibility to outsourcing (*terceirização*) based on industry reports. Again, the inclusion hardly alters trade-variable coefficients (column 7). Interestingly, our indicator for outsourcing susceptibility itself is not a significant predictor of displacements but a highly significant predictor of reduced accessions. There is, overall, no evidence that concomitant reforms overlay the impact of trade reform on labor market turnover.

Worker-fixed effects. The unconditional mean worker accession rates in sectors with comparative advantage exceed the unconditional mean rates in sectors with a comparative disadvantage (Table 5). The comparative-advantage coefficient estimate, however, exhibits the opposite negative sign in FE cLogit

(Table 9) and will turn significantly negative in the presence of higher-order interaction terms (Table 15). Why the difference? We explore the importance of worker effects and re-estimate separations and accessions with a conventional logit model. Table 14 compares estimates to those in Tables 8 and 9.

Results suggest that high-turnover workers are more frequently employed in sectors with strong comparative advantage and little import penetration (formerly protected sectors). Not conditioning on workers' turnover behavior results in a negative association of import penetration with separations and accessions, or high labor turnover in low-penetration sectors. This reverts the sign on import penetration in both the separation and accession regressions, and reverts the sign on comparative advantage in the accession regression. These insights are also consistent with the hypothesis that there is sorting of high-turnover workers into high-turnover sectors, which tend to be the sectors of comparative advantage in Brazil.

Higher-order interactions. Last, we explore interactions of trade variables and their joint effects on separations and accessions. Table 15 compares earlier separation and accessions estimates (from Tables 8 and 9) to regressions with interaction terms. There are no remarkable changes to coefficient estimates for separations. At the accession margin, however, three noteworthy changes emerge for the full set of interactions (column 6). First, the negative comparative advantage coefficient turns significant: employers in comparative-advantage sector hire workers significantly less frequently. Second, product tariff reductions depress accession rates most strongly in sectors with a comparative advantage. Third, there is the one piece of good news for classic trade theory. Although both comparative advantage sectors and exporters hire significantly fewer workers, exporters within the comparative advantage sectors partly offset these negative associations and product tariff cuts further boost the offsetting effect for exporters within comparative advantage sectors.

7 Conclusions

This paper contrasts the common finding that output shares are reallocated to more productive plants or firms after trade reform with direct evidence on the factor market. A comprehensive formal-sector worker data set allows us to tracking workers across sectors and firms in the aftermath of Brazil's large-scale trade reform. It shows that comparative-advantage sectors and exporters

Table 15: CONDITIONAL LOGIT ESTIMATES WITH INTERACTIONS

	Separations			Accessions		
	(1)	(2)	(3)	(4)	(5)	(6)
Balassa Comp. Adv.	.168	.138	.132	-.016	-.058	-.124
	(.024)***	(.036)***	(.043)***	(.020)	(.032)*	(.038)***
Comp. Adv. × Prd. Trff.		.200	.270		.289	.596
		(.200)	(.238)		(.162)*	(.203)***
Exporter Status	.283	.481	.473	-.439	-.361	-.561
	(.028)***	(.048)***	(.081)***	(.027)***	(.045)***	(.077)***
Exporter × Prd. Trff.		-1.069	-.933		-.423	.344
		(.213)***	(.362)**		(.195)**	(.323)
Comp. Adv. × Exporter			.014			.153
			(.051)			(.047)***
Comp. Adv. × Exp. × Prd. Trff.			-.155			-.672
			(.291)			(.250)***
Product Market Tariff	-.710	-.427	-.511	1.248	.966	.546
	(.426)*	(.532)	(.548)	(.393)***	(.474)**	(.504)
Intm. Input Tariff	2.893	3.249	3.299	-3.078	-2.491	-2.303
	(.678)***	(.768)***	(.767)***	(.598)***	(.672)***	(.682)***
Import Penetration	1.247	1.085	1.080	.203	.041	.005
	(.388)***	(.393)***	(.393)***	(.355)	(.364)	(.364)
Obs.	145,418	145,418	145,418	112,971	112,971	112,971
Pseudo R^2	.150	.151	.151	.041	.041	.041

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

impede, rather than foster, the formal-sector reallocations needed after trade-induced worker displacements. As a consequence, trade is associated with more frequent transitions to informal work status and unemployment, longer durations of formal-job reallocations and more frequent failures of formal-job reallocations for an extended period of time. Although product-market reallocation can be fast after trade reform, countries similar to Brazil may want to prepare for prolonged and incomplete adjustment in the labor market.

Appendix

A Data

Brazilian law requires all Brazilian plants to submit detailed annual reports with individual information on their workers and employees to the ministry of labor (*Ministério de Trabalho*, *MTE*). The collection of these reports is called *Relação Anual de Informações Sociais*, or *RAIS*, and is typically concluded at the parent firm by late February or early March for the preceding year of observation.

An plant's failure to report its workforce information can, in principle, result in fines proportional to the workforce size but fines are rarely 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 (*Abono Salarial*) is exclusively based on *RAIS* records. The ministry of labor estimates that currently 97 percent of all formally employed workers in Brazil are covered in *RAIS*, and that the coverage exceeded 90 percent throughout the 1990s.

Screening. In *RAIS*, workers are identified by individual-specific *PIS* (*Programa de Integração Social)* IDs that are similar to social security numbers in the U.S. (but *PIS* IDs are not used for identification purposes other than the administration of the wage supplement program *Abono Salarial*). A given plant may report the same *PIS* ID multiple times within a single year in order to help the worker withdraw deposits from the worker's severance pay savings account (*Fundo de Garantia do Tempo de Serviço*, *FGTS*) through spurious layoffs and rehires. Moreover, bad compliance causes certain *PIS* IDs to be recorded incorrectly or repeatedly. To handle these issues, we screen the census records as follows. (1) Observations with *PIS* IDs having fewer than 11 digits are removed. These correspond to either informal (illegal) workers or measurement error from faulty bookkeeping. (2) Multiple employments with the same accession and separation date at the same employer are removed. For a worker with such multiple employments, we only keep the observation with the highest average monthly wage level (in cases of wage level ties, we drop duplicate observations randomly).

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

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

	RAIS Age Category	Imputed Age
1.	Child (10-14)	12
2.	Youth (15-17)	16
3.	Adolescent (18-24)	21
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*

Our education variable regroups the nine education categories included in *RAIS* to correspond to five typically considered categories. We define the education indicator variables as follows:

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

The occupation indicator variables are obtained from the *CBO* classification codes in *RAIS*, as reclassified to conform with the *ISCO-88* categories.[17] The available *RAIS* version for the nation as a whole, *CBO* classes are only reported at the three-digit level. We adjust the mapping from *CBO* to *ISCO-88* accordingly. We ultimately map *ISCO-88* categories to occupation levels as follows:

[17] See our documentation online at *econ.ucsd.edu/muendler/brazil*.

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

Additional sector data. We use Ramos and Zonenschain (2000) national accounting data to calculate the *effective rate of market penetration* with foreign imports. Arguably, domestic firms find the absorption market corresponding to $A_{i,t} \equiv Y_{i,t} - (X_{i,t} - M_{i,t})$ the relevant domestic environment in which they compete. We define the effective rate of market penetration as $M_{i,t}/A_{i,t}$.

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

We use cumulated foreign direct investment stock data from the Brazilian central bank (*Banco Central do Brasil*) for 1986 through 1995. A central bank survey in 1995 suggests that cumulated FDI stocks were overestimated prior to 1995, and we correct them down by an according adjustment factor. From 1996 on, we use central bank figures of FDI flows, based on new sector definitions adopted since December 1995, to infer FDI stocks through 1998.

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

Table 16: TRADEABLE GOODS SECTORS

	Comp. adv.	Exp. ind.	Tariff Prd.	Tariff Inp.	Imp. pen.
	(1)	(2)	(3)	(4)	(5)
1 Mining & quarrying	.861	.407			
2 Manufacture of non-metallic mineral products	1.120	.288	.294	.252	.019
3 Manufacture of metallic products	1.697	.540	.228	.205	.046
4 Manufacture of machinery, equipment, instruments	.551	.615	.323	.302	.110
5 Manufacture of electrical & telecom. equipment	.576	.669	.367	.325	.168
6 Manufacture of transport equipment	1.041	.785	.458	.345	.103
7 Manufacture of wood products & furniture	1.064	.291	.228	.224	.011
8 Manufacture of pulp & paper, and publishing	.608	.386	.238	.243	.037
9 Manufacture of rubber, leather and prod. n.e.c.	.696	.593	.412	.369	.064
10 Manufacture of chemical & pharmaceutical prod.	.731	.592	.244	.198	.079
11 Manufacture of apparel & textiles	.533	.534	.470	.401	.037
12 Manufacture of footwear	3.318	.670	.328	.307	.066
13 Manufacture of food, beverages, & ethyl alcohol	3.012	.411	.273	.188	.021
25 Agriculture, hunting, forestry & fishing	1.553	.083			

Sources: UN Comtrade 1986-98; *SECEX* 1990-98 exporter status (weighted by nationwide *RAIS* jobs of prime-age male workers); product 1986-98 tariffs from Kume et al. (2000) (weighted with *IBGE* input-output matrix for input tariffs); import penetration 1986-98 from Ramos and Zonenschain (2000).

Table 17: Subsector *IBGE* and *Nível 50* Comparison

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

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

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