Firm Dynamics, Job Turnover, and Wage Distributions in an Open Economy*

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Abstract

This paper explores the combined effects of reductions in trade frictions, tariffs, and firing costs on firm dynamics, job turnover, and wage distributions. It uses establishment-level data from Colombia to estimate an open economy dynamic model that links trade to job flows and wages. Counterfactual experiments imply that Colombia's integration with global product markets increased its national income at the expense of higher job turnover and greater lifetime wage inequality. In contrast, these experiments find little role for this country's labor market reforms in driving these variables. The results speak more generally to the effects of globalization on labor markets.

Keywords: International trade, firm dynamics, size distribution, labor market frictions, inequality
JEL Codes: F12, F16, E24, J64, L11

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1 Introduction

During the late 1980s and early 1990s, as the forces of globalization gained momentum, many Latin American countries dismantled their trade barriers and implemented labor market reforms.[1] Over the next two decades, these countries roughly doubled their trade-to-GDP ratios, and thereby reaped the well-known benefits of better access to foreign markets. But they also experienced more rapid job turnover, higher unemployment rates, increased wage inequality, and greater informal sector activity.[2]

These developments motivate the two basic questions we address in this paper. First, through what mechanisms and to what extent might the global integration of product markets have increased job insecurity and wage inequality in Latin America? Second, how might commercial policy reforms and changes in worker firing costs have conditioned the relationship between globalization and these labor market outcomes?

To answer these questions, we develop a dynamic general equilibrium model that links globalization and labor regulations to job flows, unemployment, and wage distributions. Then we fit our model to plant-level panel data from Colombia—a country that cut tariffs, reduced firing costs, and exhibited rapid growth in merchandise trade. Finally, we perform counterfactual experiments that quantify the labor market consequences of global reductions in trade frictions (hereafter, "globalization") and Colombia's policy reforms. Decomposing the net effects, we find that the policy reforms modestly increased job turnover and unemployment, modestly improved average income, and left wage inequality unchanged. But globalization was much more important, accounting by itself for a substantial fraction of the increase in job turnover, unemployment, and average income that Colombia experienced. It also significantly increased cross-worker inequality in lifetime earnings. Hence, while improving average incomes through standard channels, the rapid expansion of global trade may also be contributing to reduced job security and heightened inequality in Latin America and elsewhere.[3]

Several key mechanisms in our model link openness and labor market outcomes. First, by increasing the sensitivity of firms' revenues to their productivity and employment levels, openness makes firms more willing to incur the hiring and firing costs associated with

[1] Global trade expansion stalled in the early 1980s, but resumed growing much more rapidly than production thereafter (WTO, 2014). Haltiwanger et al. (2004) document the general reduction in Latin American trade barriers during this period. Heckman and Pages (2004) survey labor market regulations and reforms in Latin America, noting that openness to international trade increased the demand for labor market flexibility.

[2] Trade data are taken from Feenstra et al. (2013). The Inter-American Development Bank (2004) summarizes the deterioration in Latin American labor market conditions. Haltiwanger et al. (2004) document the association between job turnover and openness in Latin America. Goldberg and Pavcnik (2007) survey the evidence linking openness to wage inequality and informality in Latin America and other developing countries.

[3] Rodrik (1996) makes a related argument, though he points to different mechanisms.

adjusting their workforce. By itself, this *sensitivity effect* makes job turnover and unemployment higher when trade frictions are low. It also tends to create larger rents for the more successful firms and to thereby spread the cross-firm wage distribution. Second, however, openness concentrates workers at larger firms, which are more stable than small firms and less likely to exit.[4] This *distribution effect* works against the sensitivity effect, tending to *reduce* turnover and wage inequality as trade frictions fall. In our policy experiments, the sensitivity effect proves dominant.

Our model is related to several literatures. First, it shares some basic features with large firm models in the labor-search literature. In particular, it can be viewed as an extension of Bertola and Caballero (1994), Bertola and Garibaldi (2001), and Koeniger and Prat (2007) to include fully articulated product markets, international trade, serially correlated productivity shocks, intermediate inputs, and endogenous firm entry and exit.[5] And like these models, it explains firms' wage setting as reflecting their idiosyncratic demand or productivity shocks in the presence of convex hiring costs.

Second, our work contributes to the growing literature on the effects of international trade in the presence of labor market frictions. Like many papers therein, we link the cross-firm wage distribution to the cross-firm rent distribution, which we link in turn to trade costs through a Melitz (2003) mechanism (Egger and Kreickemeier, 2009; Helpman et al., 2010 and 2012; Felbermayr et al., 2011; Fajgelbaum, 2013; Davis and Harrigan, 2011; and Amiti and Davis, 2012).[6] Among these papers, our model is relatively close to those that generate wage dispersion and unemployment by combining wage bargaining with Diamond-Mortensen-Pissarides search frictions (Helpman and Itskhoki, 2010; Helpman et al., 2010 and 2012; and Felbermayr et al., 2011). The key distinction between these papers and ours is that we empirically characterize steady-states with ongoing productivity shocks, endogenous entry and exit, and job turnover. Our model is also relatively close to Fajgelbaum's (2013), which analyzes the life cycle of firms in an open economy with on-the-job search. However, Fajgelbaum (2013) does not allow post-entry productivity shocks, and thus does not consider job turnover.

Finally, because our paper offers a new explanation for size-dependent volatility, it contributes to the literature on firm dynamics (e.g., Hopenhayn, 1992; Jovanovic, 1982; Ericson

[4] This feature of our model captures a well-known empirical regularity. Haltiwanger et al. (2013) provide recent evidence from the U.S.

[5] Other recent papers that study firm dynamics and labor market frictions in a closed economy context include Cooper et al. (2007), Lentz and Mortensen (2010), Hobijn and Sahin (2013), and Elsby and Michaels (2014). Utar (2008) studies firm dynamics and labor market frictions in an import-competing industry that takes the wage rate as given.

[6] Many other recent papers examine the effects of trade on imperfect labor markets, but presume competitive product markets and homogenous firms. Examples of theoretical and empirical work in this vein include Davidson et al. (1999, 2008), Kambourov (2009), Artuc et al. (2010), Coşar (2013), and Dix-Carneiro (2014).

and Pakes, 1995; Klette and Kortum, 2004; Luttmer, 2007; Rossi-Hansberg and Wright, 2007; and Arkolakis, 2013). Unlike previous studies, we explain the relative stability of large firms as a consequence of nonlinear hiring costs, which make it relatively more expensive (per worker) for large firms to sustain any positive growth rate.

Our estimated model closely replicates basic features of Colombian micro data in the decade preceding reforms, including the size distribution of firms, the rates of employment growth among firms of different sizes, producer entry and exit rates, exporting patterns, and the degree of persistence in firm-level employment levels. Also, although it is fit to pre-reform data, it nicely replicates many post-2000 features of the Colombian economy when it is evaluated at post-2000 tariff rates, firing costs, and global trade frictions. In particular, the quantified model successfully predicts the post-2000 plant size distribution.

While we do not pretend to capture all of the channels through which openness and firing costs can affect labor market outcomes, our focus on firm-level entry, exit and idiosyncratic productivity shocks is supported by existing empirical evidence on the sources of job turnover and wage heterogeneity. Studies of job creation and job destruction invariably find that most reallocation is due to idiosyncratic (rather than industry-wide) adjustments (Davis et al. 1998; Roberts 1996), even in Latin America's highly volatile macro environment (Chapter 2 of Inter-American Development Bank 2004). Further, as Goldberg and Pavcnik (2007) note, there is little evidence in support of trade-induced labor reallocation across sectors, so if openness has had a significant effect on job flows, it should have been through intra-sectoral effects. Finally, while observable worker characteristics do matter for wage differentials, much is attributable to labor market frictions and firm heterogeneity (Abowd et al. 1999; Mortensen 2003; Helpman et al. 2012).

2 The Model

2.1 Preferences

We consider a small open economy populated by a unit measure of homogeneous, infinitely-lived worker-consumers. Each period t, agents derive utility from the consumption of homogeneous, non-tradable services, s_t, and a composite industrial good, c_t, where

$$c_t = \left(\int_0^{N_t} c_t(n)^{\frac{\sigma-1}{\sigma}} dn \right)^{\frac{\sigma}{\sigma-1}}, \tag{1}$$

aggregates consumption of the differentiated goods varieties, $c_t(n)$, $n \in [0, N_t]$, with a constant elasticity of substitution $\sigma > 1$. Worker-consumers maximize the expected present value of their utility stream

$$\mathcal{U} = \sum_{t=1}^{\infty} \frac{s_t^{1-\gamma} c_t^{\gamma}}{(1+r)^t},$$

where r is the discount rate and $\gamma \in (0,1)$ is the expenditure share of the industrial good. Being risk neutral, they do not save. In what follows, we suppress time subscripts t for ease of notation.

2.2 Production technologies

Services are supplied by service sector firms and, less efficiently, by unemployed workers engaged in home production. Regardless of their source, services are produced with labor alone, homogeneous across suppliers, and sold in competitive product markets. Firms that supply services generate one unit of output per worker and face no hiring or firing costs. Unemployed workers who home-produce service goods each generate $b < 1$ units of output. The economy-wide supply of services is thus

$$S = L_s + bL_u, \tag{2}$$

where L_s is labor employed in the service sector and L_u is unemployed labor.

Differentiated goods are supplied by industrial sector firms, each of which produces a unique product. These firms are created through sunk capital investments; thereafter their output levels are determined by their productivity levels, z, employment levels, l, and intermediate input usage, m, according to:

$$q = z l^{\alpha} m^{1-\alpha}. \tag{3}$$

Here $0 < \alpha < 1$ and $m = \left(\int_0^N m(n)^{\frac{\sigma-1}{\sigma}} dn\right)^{\frac{\sigma}{\sigma-1}}$ aggregates differentiated goods used as intermediates in the same way the subutility function (1) aggregates differentiated goods used for final consumption. Note that, as in Melitz (2003), productivity variation can equally well be thought of as variation in product quality.

2.3 Price indices

Differentiated goods can be traded internationally. Measure N_F of the measure N differentiated goods are imported, and an endogenous set of domestically produced goods are exported. Both exports and imports are subject to iceberg trade costs: for each $\tau_c > 1$ units shipped, a single unit arrives at its destination. Moreover, imports are subject to an *ad valorem* tariff rate of $\tau_a - 1 > 0$.

Let asterisks indicate that a variable is expressed in foreign currency, and define $p^*(n)$ to be the FOB price of imported variety $n \in [0, N_F]$. The exact home-currency price index for imported goods is then $P_F = \tau_a \tau_c k \left[\int_0^{N_F} p^*(n)^{1-\sigma} dn\right]^{1/(1-\sigma)}$, where k is the exchange rate. Similarly, letting $p(n)$ be the price of domestic variety $n \in (N_F, N]$ in the home market, the exact home price index for domestic goods is $P_H = \left(\int_{N_F}^{N} p(n)^{1-\sigma} dn\right)^{1/(1-\sigma)}$. Finally, defining $p_X^*(n)$ to be the price of domestic variety n in the foreign market, and letting $\mathcal{I}^x(n) \in \{0, 1\}$ take a value of 1 if good n is exported, $P_X^* = \left(\int_{N_F}^{N} \mathcal{I}^x(n) p_X^*(n)^{1-\sigma} dn\right)^{1/(1-\sigma)}$ is the exact foreign market price index for exported goods.

Several normalizations simplify notation. First, since the measure of available foreign varieties and their FOB foreign-currency prices are exogenous to our model, we normalize $\left[\int_0^{N_F} p^*(n)^{1-\sigma} dn\right]^{1/(1-\sigma)}$ to unity by choice of foreign currency units. This allows us to write the exact domestic price index for the composite industrial good as

$$P = \left[P_H^{1-\sigma} + (\tau_a \tau_c k)^{1-\sigma}\right]^{\frac{1}{1-\sigma}}. \tag{4}$$

Second, without loss of generality, we choose the price of services to be our numeraire. The real exchange rate k endogenously adjusts so that in equilibrium, the two normalizations in domestic and foreign currency units are consistent.

2.4 Differentiated goods markets

Differentiated goods are sold in monopolistically competitive markets, where they are purchased by consumers as final goods and by producers as intermediate inputs. Utility maximization implies that worker i with income Y_i demands $\frac{\gamma Y_i}{P}\left(\frac{p(n)}{P}\right)^{-\sigma}$ units of domestic variety n and $\frac{\gamma Y_i}{P}\left(\frac{\tau_a \tau_c k p^*(n')}{P}\right)^{-\sigma}$ units of imported variety n'. Similarly, firm j with gross revenue G_j optimally purchases $(1-\alpha)\frac{\sigma-1}{\sigma}\frac{G_j}{P}\left(\frac{p(n)}{P}\right)^{-\sigma}$ units of domestic variety n, and $(1-\alpha)\frac{\sigma-1}{\sigma}\frac{G_j}{P}\left(\frac{\tau_a \tau_c k p^*(n')}{P}\right)^{-\sigma}$ units of imported variety n'.

Aggregating across domestic consumers and domestic producers yields total domestic demand for any domestic variety n:

$$Q_H(n) = D_H p(n)^{-\sigma} \quad \text{for} \quad n \in (N_F, N], \tag{5}$$

where

$$D_H = P^{\sigma-1}\left[\gamma \int_0^1 Y_i di + (1-\alpha)\frac{\sigma-1}{\sigma}\int_{N_F}^{N} G_j dj\right].$$

Note that the population of domestic worker-consumers is normalized to one, and domestic producers are indexed by $n \in (N_F, N]$. Likewise, total domestic demand for any imported variety n is

$$Q_H(n) = D_H \left[\tau_a \tau_c k p^*(n)\right]^{-\sigma} \quad \text{for} \quad n \in [0, N_F]. \tag{6}$$

Finally, assuming markets are internationally segmented, foreign demand for domestically produced good n is given by

$$Q_F(n) = D_F^* \left[p_X^*(n)\right]^{-\sigma}, \; n \in (N_F, N], \tag{7}$$

where D_F^* measures aggregate expenditures abroad denominated in foreign currency, and is net of any effects of foreign commercial policy. Given our small country assumption, we take D_F^* to be unaffected by the actions of domestic agents.

These expressions imply that, expressed in domestic currency, total domestic expenditures on domestic varieties amount to $D_H P_H^{1-\sigma}$, total domestic expenditures on imported varieties amount to $D_H \left(\tau_a \tau_c k\right)^{1-\sigma}$, and domestic firms' total export revenues amount to $k D_F^* P_X^{*1-\sigma} / \tau_c$.

2.5 Producer dynamics

Industrial firms are subject to idiosyncratic productivity shocks. These are generated by the $AR(1)$ process

$$\ln z' = \rho \ln z + \sigma_z \epsilon, \tag{8}$$

where $\rho \in (0, 1)$ and $\sigma_z > 0$ are parameters, primes indicate one-period leads, and $\epsilon \sim N(0, 1)$ is a standard normal random variable independently and identically distributed across time and firms. Together with firms' employment policies and entry/exit decisions, (8) determines the steady state distribution of firms over the state space (z, l) and the rates at which firms transit across pairs of states.

Producer dynamics in the industrial sector resemble those in Hopenhayn (1992) and Hopenhayn and Rogerson (1993) in that firms react to their productivity shocks by optimally hiring, firing, or exiting. Also, new firms enter whenever their expected future profit stream exceeds the entry costs they face. However, unlike these papers, we assume that hiring in the industrial sector is subject to search frictions captured by a standard matching function. We now describe the functioning of labor markets.

2.6 Labor markets and the matching technology

The service sector labor market is frictionless, so workers can obtain jobs there with certainty if they choose to do so. Since each service sector worker produces one unit of output, and the price of services is our numeraire, these jobs pay a wage of $w_s = 1$.

The industrial sector labor market, in contrast, is subject to search frictions. These expose industrial job seekers to unemployment risk and create match-specific rents that workers and firms bargain over. The number of new matches between job seekers and vacancy posting firms each period is given by

$$M(V,U) = \frac{VU}{(V^\theta + U^\theta)^{1/\theta}},$$

where $\theta > 0$. Here, U is the measure of workers searching for industrial sector jobs, and V is the measure of industrial sector vacancies. The parameter θ governs the severity of matching frictions, since a higher value for θ results in a larger number of matches for given values of U and V.[7] This matching function implies that industrial firms fill each vacancy with probability

$$\phi(V,U) = \frac{M(V,U)}{V} = \frac{U}{(V^\theta + U^\theta)^{1/\theta}},$$

while workers searching for industrial jobs find matches with probability

$$\widetilde{\phi}(V,U) = \frac{M(V,U)}{U} = \frac{V}{(V^\theta + U^\theta)^{1/\theta}}.$$

At the beginning of each period, workers who are not already employed in the industrial sector decide whether to accept a service sector job that pays wage $w_s = 1$ with certainty, or to search for an industrial sector job. If they fail to match with an industrial sector producer, they subsist until the next period by home-producing services at the wage of $b < 1$.[8] At the start of the matching process, among the unit measure of the worker population, U are searching for an industrial job. At the end of the matching process, $L_u = (1 - \widetilde{\phi})U$ workers fail to find a job and stay unemployed while L_q work in the industrial sector. As a result, a fraction $L_u/(L_u + L_q)$ of workers associated with the frictional labor market are unemployed.

Workers who begin a period employed in the industrial sector can continue with their current job unless their employer lays them off or shuts down entirely. In equilibrium, industrial sector workers are paid at least their reservation wage, so those who do not lose their jobs will never leave them voluntarily. Workers' job-seeking decisions and the bargaining game that determines industrial firms' wages will be described below in Sections 2.9 and 2.10, respectively. But before discussing either, we must characterize the firm's problem.

[7] The functional form of the matching function follows den Haan et al. (2000). It is subject to constant returns to scale, and increasing in both arguments. In contrast to the standard Cobb-Douglas form, it has no scale parameter and the implied matching rates are bounded between zero and one. Note that for $V = U$, as θ approaches infinity, job finding and filling probabilities approach to one.

[8] The notion that workers trade job security in a low wage sector for the opportunity to search in a higher wage sector traces back at least to the Harris and Todaro (1970) model.

2.7 The firm's problem

At the beginning of each period, incumbent firms decide whether to continue operating and potential entrants decide whether to create new firms. Thereafter, active firms go on to choose their employment levels, intermediate input usage, and exporting policies. Entry, exit, hiring, and firing involve adjustment costs, so these decisions are solutions to forward-looking problems. In contrast, intermediate input purchases and exporting decisions involve frictionless optimization after employment levels have been determined. We now characterize all firm decisions.

2.7.1 Export policy

Given the domestic demand function (5), any firm that sells some fraction $1-\eta$ of its output domestically will generate gross home sales amounting to $D_H^{\frac{1}{\sigma}}\left[(1-\eta)q\right]^{\left(\frac{\sigma-1}{\sigma}\right)}$. Similarly, given the foreign demand function (7), such a firm will generate gross foreign sales of $k\left(D_F^*\right)^{\frac{1}{\sigma}}\left[\frac{\eta}{\tau_c}q\right]^{\frac{\sigma-1}{\sigma}}$. Total gross revenue can thus be written as

$$G(q,\eta)=\exp\left[d_H+d_F(\eta)\right]q^{\frac{\sigma-1}{\sigma}}, \tag{9}$$

where $d_H=\ln(D_H^{\frac{1}{\sigma}})$ and $d_F(\eta)=\ln\left[(1-\eta)^{\frac{\sigma-1}{\sigma}}+\ k\left(\frac{D_F^*}{D_H}\right)^{\frac{1}{\sigma}}(\eta/\tau_c)^{\frac{\sigma-1}{\sigma}}\right]$. While the term d_H measures domestic demand, and is common to all firms, the term $d_F(\eta)$ captures the extra revenue generated by exporting, conditional on output.

Given output levels, firms choose their exporting levels each period to maximize their current sales revenues net of fixed exporting costs, c_x. Not all firms find it profitable to participate in foreign markets, but those that do share the same optimal level of η:

$$\eta^o=\arg\max_{0\le\eta\le1}d_F(\eta)=\left(1+\frac{\tau_c^{\sigma-1}D_H}{k^\sigma D_F^*}\right)^{-1}. \tag{10}$$

The associated export market participation policy is thus

$$\mathcal{I}^x(q)=\begin{cases}1 & \text{if }\left[\exp\left[d_H+d_F(\eta^o)\right]-\exp(d_H)\right]q^{\frac{\sigma-1}{\sigma}}>c_x,\\ 0, & \text{otherwise,}\end{cases} \tag{11}$$

and there is a threshold output level that separates exporters from others. Given d_H and $d_F(\eta^o)$, this allows us to write revenues net of exporting costs as a function of output alone:

$$G(q)=\exp\left[d_H+\mathcal{I}^x(q)d_F(\eta^o)\right]q^{\frac{\sigma-1}{\sigma}}-c_x\mathcal{I}^x(q). \tag{12}$$

2.7.2 Intermediates and the value-added function

Firms determine their output levels by choosing their intermediate input usage, m, given their current period z and l values. Optimizing over m and suppressing market-wide variables, we can thus use (3) and (12) to write value added net of exporting costs as a function of z and l alone:

$$R(z,l) = \max_{m} \left\{ G(zl^{\alpha}m^{1-\alpha}) - Pm \right\}. \tag{13}$$

The solution to this optimization problem is:

$$R(z,l) = \Delta(z,l)\,(zl^{\alpha})^{\Lambda} - c_x \mathcal{I}^x(z,l), \tag{14}$$

where we have used the optimized m value and (3) to restate $\mathcal{I}^x(q)$ as $\mathcal{I}^x(z,l)$. Also,

$$\Delta(z,l) = \Theta P^{-(1-\alpha)\Lambda} \Bigg(\exp\left[d_H + \mathcal{I}^x(z,l) d_F(\eta^o)\right] \Bigg)^{\frac{\sigma}{\sigma-1}\Lambda}, \tag{15}$$

where $\Lambda = \frac{\sigma-1}{\sigma-(1-\alpha)(\sigma-1)}$ and $\Theta = \left(\frac{1}{(1-\alpha)\Lambda}\right)\left[\frac{(1-\alpha)(\sigma-1)}{\sigma}\right]^{\frac{\sigma}{\sigma-1}\Lambda}$ are positive constants.

The term $\Delta(z,l)$ is a firm-level market size index. It responds to anything that affects aggregate domestic demand (D_H), trade costs (τ_c), or the exchange rate (k). But given these market-wide variables, the only source of cross-firm variation in $\Delta(z,l)$ is exporting status ($\mathcal{I}^x$). Accordingly, below we suppress the arguments of Δ except where we wish to emphasize its dependence on these variables. Appendix 1 provides derivations of (14) and (15), and shows that the net revenue function exhibits diminishing marginal returns to labor ($\alpha\Lambda < 1$).

2.7.3 Employment policy

We now turn to decisions that involve forward-looking behavior. When choosing employment levels, firms weigh the revenue stream implied by (14) against wage costs, the effects of l on their continuation value, and current firing or hiring costs. To characterize the latter, let the cost of posting v vacancies for a firm of size l be

$$C_h(l,v) = \left(\frac{c_h}{\lambda_1}\right)\left(\frac{v}{l^{\lambda_2}}\right)^{\lambda_1},$$

where c_h and $\lambda_1 > 1$ are positive parameters.[9] The parameter $\lambda_2 \in [0,1]$ determines the strength of scale economies in hiring. If $\lambda_2 = 0$, there are no economies of scale and the cost

[9] This specification generalizes Nilsen et al. (2007), who set $\lambda_2 = 1 - 1/\lambda_1$. See also Merz and Yashiv (2007), and Yashiv (2006).

of posting v vacancies is the same for all firms. On the other hand, if $\lambda_2 = 1$, the cost of a given employment growth rate is the same for all firms.

All firms in our model are large enough that cross-firm variation in realized worker arrival rates is ignorable. That is, all firms fill the same fraction ϕ of their posted vacancies. It follows that expansion from l to l' simply requires the posting of $v = \frac{l'-l}{\phi}$ vacancies, and we can write the cost of expanding from l to l' workers as

$$C_h(l, l') = \left(\frac{c_h}{\lambda_1}\right) \phi^{-\lambda_1} \left(\frac{l' - l}{l^{\lambda_2}}\right)^{\lambda_1}. \tag{16}$$

Clearly, when labor markets are slack, hiring is less costly because each vacancy is more likely to be filled. Also, for $\lambda_1(1-\lambda_2) > 1$, a given level of employment growth is more costly per worker for larger firms.[10] So, other things equal, $\lambda_1(1-\lambda_2) > 1$ means that larger firms expand relatively slowly in response to positive shocks, and they have stronger incentives to hoard labor in the face of transitory negative shocks. This feature of our model is the main reason it is able to replicate the well-known association between firm size and job stability.

Downward employment adjustments are also costly. When a firm reduces its workforce from l' to l, it incurs firing costs proportional to the number of workers shed:[11]

$$C_f(l, l') = c_f(l - l'). \tag{17}$$

For convenience we assume hiring and firing costs are incurred in terms of service goods, and we describe both with the adjustment cost function:

$$C(l, l') = \begin{cases} C_h(l, l') & \text{if } l' > l, \\ C_f(l, l') & \text{otherwise.} \end{cases}$$

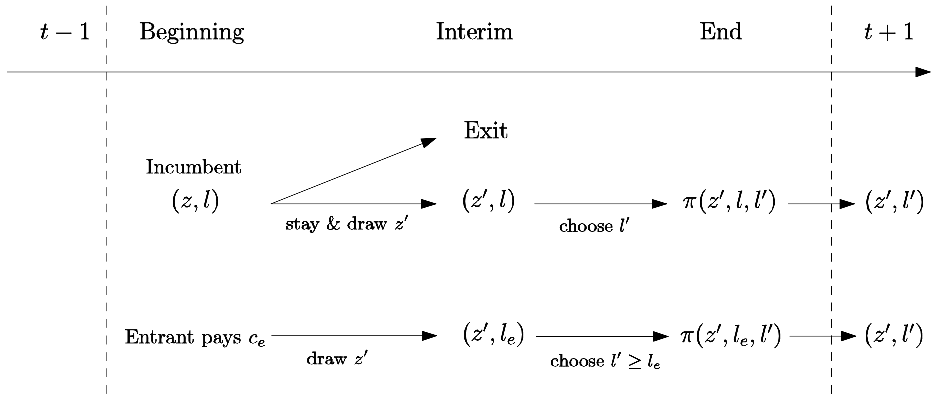
Several observations concerning adjustment costs are in order. First, while convex hiring costs induce firms to expand gradually, firms cannot economize on firing costs by downsizing gradually. Second, when the firm exits, it is not liable for c_f. Finally, as will be discussed below, it is possible that a firm will find itself in a position where the marginal worker reduces operating profits, but it is more costly to fire her than retain her.

Regardless of whether a firm expands, contracts, or remains at the same employment level, we assume it bargains with each of its workers individually and continuously. This implies that bargaining is over the marginal product of labor, and all workers at a firm in a

[10] By (16), labor force expansion at any given rate $g = \frac{l'-l}{l}$ involves vacancy posting costs $\left(\frac{c_h}{\lambda_1}\right)\phi^{-\lambda_1} g^{\lambda_1} l^{\lambda_1(1-\lambda_2)}$. So for $\lambda_1(1-\lambda_2) > 1$, $C_h(l, l')/l$ is increasing in l.

[11] As it is standard in the literature (see Ljungqvist 2002 for a review), we assume that firing costs take the form of a resource cost and are not pure transfers from firms to workers.

Figure 1: **Within-period Sequencing of Events for Firms**



particular state (z, l) are paid the same wage (Stole and Zwiebel 1996; Cahuc and Wasmer 2001; Cahuc, Marque, and Wasmer 2008). Moreover, the marginal worker at an expanding firm generates rents, while the marginal worker at a contracting firm does not (Bertola and Caballero 1994; Bertola and Garibaldi 2001; Koeniger and Prat 2007). Hence expanding firms face different wage schedules than others. These schedules depend upon firms' states, so we denote the wage schedule paid by a hiring firm as $w_h(z, l)$ and the wage schedule paid by a non-hiring firm as $w_f(z, l)$. Details are deferred to Section 2.10 below.

We now elaborate firms' optimal employment policies within a period (see Figure 1). An incumbent firm enters the current period with the productivity level and work force (z, l) determined in the previous period. Thereupon it may exit immediately, either because the expected present value of its profit stream is negative, or because it is hit with an exogenous exit shock.

If a firm opts to stay active and is not hit with an exogenous exit shock, it proceeds to an *interim stage* in which it observes its current-period productivity realization z'. Then, taking stock of its updated state, (z', l), the relevant wage schedules, and adjustment costs $C(l, l')$, it chooses its current period work force, l'. Both hiring and firing decisions take immediate effect and firms enter the end of the period with (z', l'), making optimal intermediate usage and exporting decisions based on their new state. Profits are realized and wages are paid at this point. Depending on whether the firm is hiring or not, profits are

$$\pi(z', l, l') = \begin{cases} R(z', l') - w_h(z', l')l' - C(l, l') - c_p & \text{if } l' > l \\ R(z', l') - w_f(z', l')l' - C(l, l') - c_p & \text{otherwise,} \end{cases} \tag{18}$$

where c_p, the per-period fixed cost of operation, is common to all firms.

Firms discount the future at the same rate $(1 + r)$ as consumers. So the beginning-of-period value of a firm in state (z, l) is

$$\mathcal{V}(z,l) = \max\left\{0, \quad \frac{1-\delta}{1+r} E_{z'|z} \max_{l'} \left[\pi(z', l, l') + \mathcal{V}(z', l')\right]\right\}, \tag{19}$$

where δ is the probability of an exogenous death shock, and the maximum of the term in square brackets is the value of the firm in the interim state, after it has realized its productivity shock. All payoffs are discounted at the interim period.

The solution to (19) implies an employment policy function,

$$l' = L(z', l), \tag{20}$$

an indicator function $\mathcal{I}^h(z', l)$ that distinguishes hiring and firing firms, and an indicator function $\mathcal{I}^c(z, l)$ that characterizes firms' continuation and exit policy. $\mathcal{I}^h(z', l)$ and $\mathcal{I}^c(z, l)$ take the value one if a firm is hiring or continuing, respectively, and zero otherwise.

2.7.4 Entry

In the steady state, a constant fraction of firms exits the industry either endogenously or exogenously. These firms are replaced by an equal number of entrants, who find it optimal to pay a sunk entry cost of c_e and create new firms. Upon entry, these entrants are endowed with an initial employment of $l_e > 0$, and draw their initial productivity level from the ergodic productivity distribution implied by (8), hereafter denoted as $\psi_e(z)$. Firms need at least l_e workers to operate, and the search costs for the initial l_e workers are included in c_e, along with fixed capital costs. Thereafter entrants behave exactly like incumbent firms, with their interim state given by (z, l_e) (see Figure 1). So by the time they begin producing, new entrants have adjusted their workforce to $l' \geq l_e$ in accordance with their initial productivity. Free entry implies that

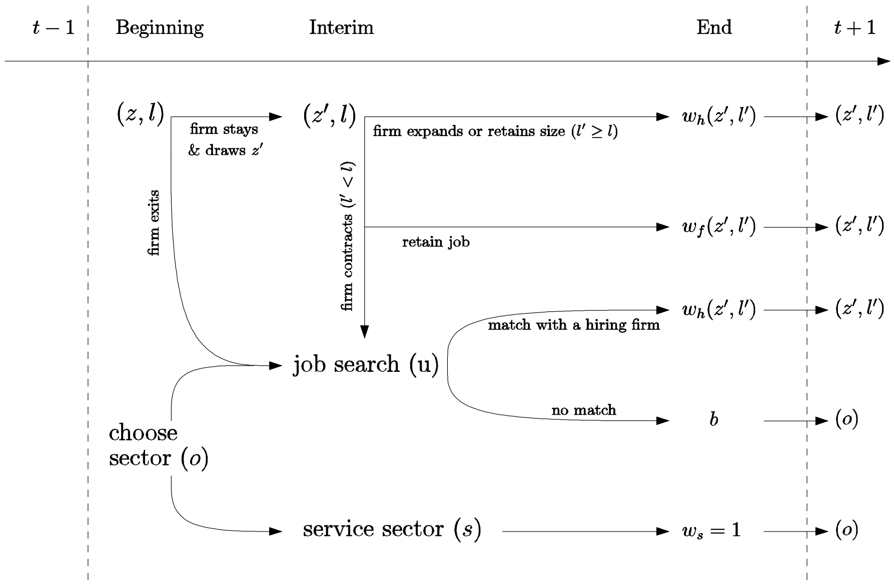
$$\mathcal{V}_e = \int_z \mathcal{V}(z, l_e)\psi_e(z)dz \leq c_e, \tag{21}$$

which holds with equality if there is a positive mass of entrants. We assume that each worker-consumer owns equal shares in a diversified fund that collects profits from firms, finances entry, and redistributes the residual as dividends to its owners.

2.8 The worker's problem

Figure 2 presents the intra-period timing of events for workers. Consider first a worker who is employed by an industrial firm in state (z, l) at the beginning of the current period. This worker learns immediately whether her firm will continue operating. If it shuts down, she joins the pool of industrial job seekers (enters state u) in the interim stage. Otherwise,

Figure 2: **Within-period Sequencing of Events for Workers**



she enters the interim stage as an employee of the same firm she worked for in the previous period. Her firm then realizes its new productivity level z' and enters the interim state (z', l). At this point her firm decides whether to hire workers. If it expands its workforce to $l' > l$, she earns $w_h(z', l')$, and she is positioned to start the next period at a firm in state (z', l'). If the firm contracts or remains at the same employment level, she either loses her job and reverts to state u or she retains her job, earns $w_f(z', l')$, and starts the next period at a firm in state (z', l'). All workers at contracting firms are equally likely to be laid off, so each loses her job with probability $p_f = (l - l')/l$.

Workers in state u are searching for industrial jobs. They are hired by entering and expanding firms that post vacancies. If they are matched with a firm, they receive the same wage as those who were already employed by the firm. If they are not matched, they support themselves by home-producing $b < 1$ units of the service good. At the start of the next period, they can choose to work in the service sector (enter state s) or search for a job in the industrial sector (remain in state u). Likewise, workers who start the current period in the service sector choose between continuing to work at the service wage $w_s = 1$ and entering the pool of industrial job seekers. These workers are said to be in state o.

We now specify the value functions for the workers in the interim stage. Going into the service sector generates an end-of-period income of 1 and returns a worker to the o state at the beginning of the next period. Accordingly, the interim value of this choice is

$$J^s = \frac{1}{1+r}(1 + J^o). \tag{22}$$

Searching in the industrial sector exposes workers to the risk of spending the period unemployed, and supporting themselves by home-producing b units of the service good. But it also opens the possibility of landing in a high-value job. Since the probability of finding a match is $\widetilde{\phi}$, the interim value of searching for an industrial job is

$$J^u = \left[\widetilde{\phi} E J_h^e + \frac{(1-\widetilde{\phi})}{1+r} (b + J^o) \right], \tag{23}$$

where EJ_h^e is the expected value of matching with a hiring firm to be defined below.

The value of the sectorial choice is $J^o = \max\{J^s, J^u\}$. In an equilibrium with both sectors in operation, workers must be indifferent between them, so $J^o = J^s = J^u$. Combined with (22), this condition implies that J^o, J^s, and J^u are all equal to $1/r$.

The expected value of matching with an industrial job, EJ_h^e, depends on the distribution of hiring firms and the value of the jobs they offer. For workers who match with a hiring firm in the interim state (z', l), the interim period value is given by

$$J_h^e(z', l) = \frac{1}{1+r} \left[w_h(z', l') + J^e(z', l') \right], \tag{24}$$

where $l' = L(z', l)$ and $J^e(z', l')$ is the value of being employed at an industrial firm in state (z', l') at the start of the next period. Accordingly, the expected value of a match for a worker as perceived at the interim stage is

$$EJ_h^e = \int_{z'} \int_l J_h^e(z', l) g(z', l) dl dz', \tag{25}$$

where $g(z', l)$ is the density of vacancies across hiring firms

$$g(z', l) = \frac{v(z', l)\widetilde{\psi}(z', l)}{\int_{z'} \int_l v(z', l)\widetilde{\psi}(z', l) dl dz'}. \tag{26}$$

Here $v(z', l) = \mathcal{I}^h(z', l) \left[L(z', l) - l \right] / \phi$ gives the number of vacancies posted by a firm in interim state (z', l), and $\widetilde{\psi}(z', l)$ is the interim stage unconditional density of firms over (z', l). Note that the latter density is distinct from the end-of-period stationary distribution of firms, $\psi(z, l)$.

It remains to specify the value of starting the period matched with an industrial firm, $J^e(z, l)$, which appears in (24) above. The value of being at a firm that exits immediately (exogenously or endogenously) is simply the value of being unemployed, J^u. This is also the value of being at a non-hiring firm, since workers at these firms are indifferent between being fired and retained. Hence $J^e(z,l)$ can be written as

$$J^e(z,l) = \left[\delta + (1-\delta)\big(1-\mathcal{I}^c(z,l)\big)\right] J^u \\ + (1-\delta)\mathcal{I}^c(z,l)\max\left\{J^u, E_{z'|z}\left[\mathcal{I}^h(z',l)J^e_h(z',l) + (1-\mathcal{I}^h(z',l))J^u\right]\right\}. \quad (27)$$

2.9 Wage schedules

We now characterize the wage schedules. Consider first a hiring firm. After vacancies have been posted and matching has taken place, the labor market closes. Firms then bargain with their workers simultaneously and on a one-to-one basis, treating each worker as the marginal one. At this point, vacancy posting costs are already sunk and workers who walk away from the bargaining table cannot be replaced in the current period. Similarly, if an agreement between the firm and the worker is not reached, the worker remains unemployed in the current period. These timing assumptions create rents to be split between the firm and the worker.

As detailed in Appendix 2, it follows that the wage schedule for hiring firms with an end-of-period state (z', l') is given by

$$w_h(z',l') = (1-\beta)b + \frac{\beta}{1-\beta+\alpha\beta\Lambda}\underbrace{\Delta(z',l')\alpha\Lambda\,(z')^{\Lambda}\,(l')^{\alpha\Lambda-1}}_{=\partial R(z',l')/\partial l'} - \beta P_f(z',l')c_f, \quad (28)$$

where $\beta \in [0,1]$ measures the bargaining power of the worker, and $P_f(z', l')$ is the probability of being fired next period.[12] Workers in expanding firms get their share of the marginal product of labor plus $(1-\beta)$ share of their outside option, while part of the firing cost is passed on to them as lower wages.

The marginal worker at a non-hiring firm generates no rents, so the firing wage just matches her reservation value (see Appendix 2):

$$w_f(z',l') = rJ^u - [J^e(z',l') - J^u]. \quad (29)$$

Three assumptions lie behind this formulation. First, workers who quit do not trigger firing costs for their employers. Second, firms cannot use mixed strategies when bargaining with workers. Finally, fired workers are randomly chosen. The first assumption ensures that workers at contracting firms are paid no more than the reservation wage, and the remaining assumptions prevent firms from avoiding firing costs by paying *less* than reservation wages to those workers they wish to shed. Importantly, $w_f(z,l)$ does vary across firing firms, since workers who continue with such firms may enjoy higher wages in the next period. This

[12] This expression is analogous to equation (9) in Koeniger and Prat (2007).

option to continue has a positive value, captured by the bracketed term in (29), so firing firms may pay their workers less than the flow value of being unemployed.

2.10 Equilibrium

Six basic conditions characterize the equilibrium. First, the distribution of firms over (z, l) states in the interim and at the end of each period, denoted by $\widetilde{\psi}(z, l)$ and $\psi(z, l)$, respectively, reproduce themselves each period through the stochastic process on z, the policy functions, and the productivity draws that firms receive upon entry. Second, all markets clear: supply matches demand for services and for each differentiated good, where supplies are determined by employment and productivity levels in each firm. Third, the flow of workers into unemployment matches the flow of workers out of unemployment—that is, the Beveridge condition holds. Fourth, a positive mass of entrants replaces exiting firms every period so that free entry condition (21) holds with equality. Fifth, aggregate income matches aggregate expenditure, so trade is balanced. Finally, workers optimally choose the sector in which they are working or seeking work. Appendix 3 provides further details.

2.11 Discussion

Firm-level patterns Before moving on to quantitative analysis, it is instructive to discuss how our model generates variation in wages and job turnover across firms in different states. To begin, consider what would happen if, at some point in time, we were to set $\rho = 1$ and $\sigma_z = 0$, making the current set of productivity draws permanent. Thereafter, given search frictions and convex hiring costs, those firms with sufficiently high draws would gradually add workers, and they would exhibit relatively high wages as they expanded toward their long run desired size. (Since expanding firms have high marginal revenue products of labor, they have more surplus to share with their workers through the bargaining game.[13]) And as they approach their long run desired sizes, at which the marginal worker generates no rents, their workers will begin to be paid their reservation wages. On the other hand, those firms with low draws that wish to contract would pay reservation wages to their workers for all periods thereafter, since linear firing costs would induce them to immediately contract to their long run size, at which workers generate no rents.

Without any mean reversion in the productivity process (i.e., with $\rho \geq 1$), expanding firms would tend to be larger than contracting firms, and the association between expansion

[13] This feature of the model is not specific to the random search framework: Felbermayr et al. (2014) show that convex vacancy posting costs play a similar role in generating residual wage inequality in a directed search model. This is in contrast to Felbermayr et al. (2011) where vacancy posting costs are linear and independent of size. Firms then immediately expand to equalize the marginal revenue product of labor to the expected marginal recruitment cost. The latter being equal across all firms, there is no wage inequality.

and high wages we have just described would induce a positive cross-sectional correlation between wages and firm size. But a sufficiently strong tendency toward mean reversion could flip this relationship by making positive productivity shocks more likely to occur at small firms, and negative shocks more likely to occur at large ones. The size-wage relationship is further complicated by the fact that vacancy posting costs per worker (16) vary across firm sizes. This makes the speed with which the rents from a positive shock dissipate depend upon firm size.

Given all of these mechanisms, each of which depends upon parameter values, it is not possible to make general statements about the size-wage relationship in our setting. Nonetheless, as Bertola and Garibaldi (2001) demonstrate, this class of model is capable of generating a positive cross-sectional correlation between employment and wages.

What about the relationship between wages and exporting? Firms expanding into export markets enjoy a discrete jump in revenues as they reallocate their sales away from domestic consumers, driving up their prices and surplus in the process. Thus, our model is consistent with the common finding that exporters tend to pay their workers better (e.g., Bernard and Jensen 1995), and it falls into the "wider class of models in which wages are increasing in firm revenue and there is selection into export markets" (Helpman et al., 2010, p. 1240).[14] However, unlike others in this class, our formulation implies exporters will eventually expand enough to drive the value of a marginal worker to zero. It thus delivers a distinctive explanation for the finding that long term exporters do not routinely pay higher wages than non-exporters (Bernard and Jensen, 1999).

Implications for exporters' mark-ups and productivity also obtain. First, the higher average prices charged by new exporters explain de Loecker and Warzynski's (2012) finding of higher mark-ups among Slovenian exporting firms. Frictionless trade models with CES preferences cannot explain this result because firms in these models freely expand or contract until their mark-ups are the same. Second, since exporters' higher prices translate into higher revenue per unit input bundle, our model provides a pricing-based explanation for the common finding that exporters enjoy higher revenue productivity.

[14] Within this class of model, size and profitability are generally associated with exporting for standard Melitz-type (2003) reasons. But wages are linked to these firm characteristics through a variety of mechanisms. For example, in Helpman et al. (2010), firms screen workers to improve their average productivity, and incentives to do so increase with firms' sales volume. The exporter wage premium then follows from the fact that highly screened workers command the highest wages and are concentrated at exporters. In Fajgelbaum's (2013) on-the-job search model, relatively productive firms expand into export markets by poaching workers from other firms, and they pay relatively high wages in order to do so. In the "fair wage" models of Amiti and Davis (2008) and Egger and Kreickmeier (2009), firms must share their rents with their workers to keep them from shirking.

Aggregates and openness Let us turn now to aggregate outcomes and their relation to increased openness. Note that our value-added function (14) links job turnover to the firm-level market size index, Δ (expression 15). Hiring cost functions play role here as well. Other things equal, an increase in Δ makes the value-added function (14) steeper, increasing the cost of deviating from static profit-maximizing employment levels. Thus, holding the distribution of firms over (z, l) fixed, policies that increase Δ will make firms' employment levels more responsive to z shocks.[15] This effect of openness on turnover works through a Melitz-type (2003) mechanism. Specifically, a reduction in τ_c or τ_a reduces Δ for unproductive non-exporters because these firms experience increased import competition without any offsetting increase in foreign sales. Exporters, on the other hand, see an increase in Δ as trade costs fall because they gain more in foreign demand than they lose in domestic demand. And small, productive non-exporters are similarly affected, since a large Δ value for exporters creates strong incentives for them to hire and expand into foreign markets. These firms therefore grow faster when exporters' Δ increases, especially when they become exporters and their Δ value jumps discretely.

This tendency for openness to make productive firms more responsive to z shocks is what we dubbed the sensitivity effect in the introduction.[16] When the heightened volatility among productive firms dominates reductions in volatility at unproductive firms, it increases job turnover. Further, to the extent that turnover rises, it is amplified by a feedback effect. Greater job turnover increases the pool of unemployed workers, and increases the vacancy filling rate, $\phi(V, U)$. This flattens the marginal cost of hiring, making firms even more responsive to z shocks.

Thus far we have discussed the effects of openness on firms' hiring policies, given their initial states, (z, l). But as Melitz (2003) stresses, the distribution of firms over the state space also reacts to changes in τ_c or τ_a. And since reductions in trade frictions tend to concentrate workers at large, stable firms, this distribution effect by itself creates a direct relationship between openness and job stability. The net effect of reductions in τ_c and τ_a

[15] To better understand this feature of our model, suppose the marginal value of an additional worker is simply her marginal revenue product, $\alpha\Lambda\Delta z\,(l')^{\alpha\Lambda-1}$, and assume the entire cost of hiring l' workers is captured by the vacancy posting cost, $\left(\frac{c_h}{\lambda_1}\right)\phi^{-\lambda_1}\frac{(l'-l)^{\lambda_1}}{l^{\lambda_2}}$. Then the first order condition for employment implies a positive relationship between l' and Δ among all firms in states where hiring occurs: $l' = f(\Delta|z, l)$, $f_\Delta > 0$. Further, the elasticity of l' with respect to z increases with Δ :

$$\frac{d\ln l'}{d\ln z} = \left[\frac{(\lambda_1 - 1)\, f(\Delta|z, l)}{f(\Delta|z, l) - l} + 1 - \alpha\Lambda\right]^{-1}.$$

Of course, other properties of our model complicate this relationship, including wage schedules, firing costs, and the distinction between the value of a worker and her marginal revenue product.

[16] Holmes and Stevens (2013) document that large firms in the U.S. are more sensitive to import competition than small firms. Their interpretation, however, differs from ours.

on turnover and unemployment thus depends upon the relative strengths of the sensitivity effect and the distribution effect.

Similarly, wage inequality responds to reductions in τ_c or τ_a through a direct effect and a distribution effect. Lower trade costs increase the additional surplus obtained by exporting, thereby increasing the exporter wage premium. In contrast, small and unproductive non-exporters face more import competition and thus have less surplus to bargain, which reduces their wages. This spreads out the cross-firm wage schedules observed in equilibrium. Through the distribution effect, however, more workers are allocated to larger firms, which tends to reduce cross-worker wage inequality. The aggregate effect once again depends on the relative strengths of these counteracting channels.

3 Quantitative Analysis

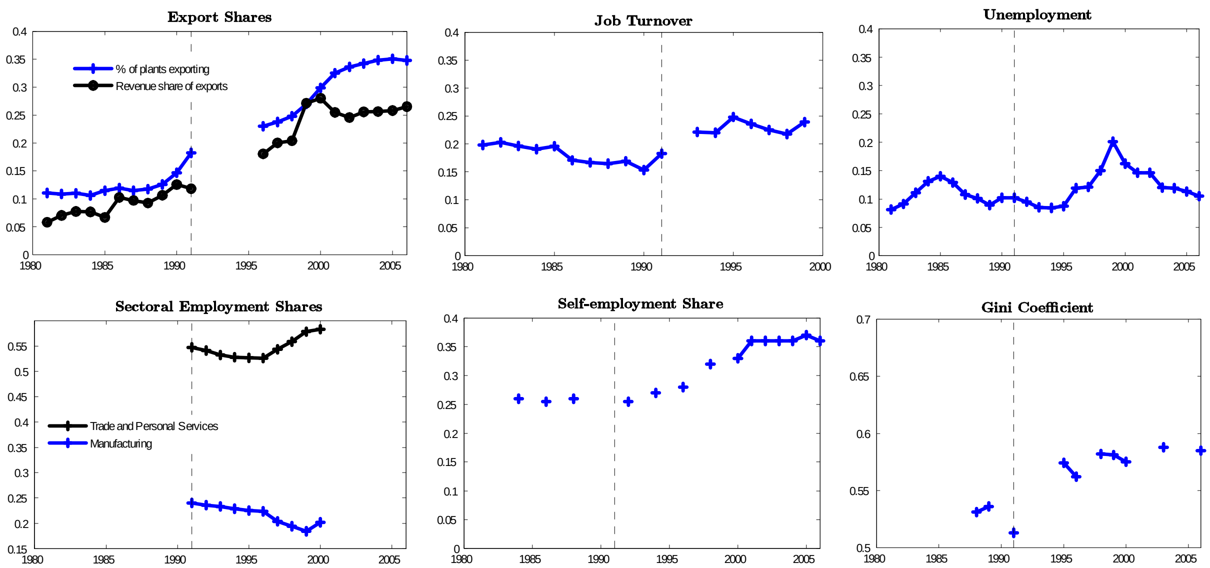
3.1 Pre- and post-reform conditions in Colombia

To explore the quantitative implications of our model, we fit it to Colombian data. This country suits our purposes for several reasons. First, Colombia underwent a significant trade liberalization during the late 1980s and early 1990s, reducing its average nominal tariff rate from 21 percent to 11 percent (Goldberg and Pavcnik 2004). Second, Colombia also implemented labor market reforms in 1991 that substantially reduced firing costs. According to Heckman and Pages (2000), the average cost of dismissing a worker fell from an equivalent of 6 months' wages in 1990 to 3 months' wages in 1999. Finally, major changes in Colombian trade volumes and labor markets followed these reforms, suggesting that they and/or external reductions in trade frictions may well have been important.

Key features of the Colombian economy during the pre- and post-reform period are summarized in Figure 3. The first panel shows the fraction of manufacturing establishments that were exporters, as well as the aggregate revenue share of exports. Before 1991, about 12 percent of all plants were exporters on average, and total exports accounted for 9 percent of aggregate manufacturing revenues. Reflecting the globalization of the Colombian economy, both ratios increased by about 250 percent from the 1980s to the 2000s. The second panel shows manufacturing job turnover due to entry, exit, and changing employment levels among continuing producers. This series went from an average of 18.1 percent during the pre-reform period 1981-1990 to 23 percent during the post-reform period 1993-1998.

The third panel of Figure 3 shows the evolution of the urban unemployment rate. During the post-reform years 1991-1998, this series hovered around its 1981-90 average of 10.8 percent. During 2000-2006, its average was a somewhat higher 13 percent, but this increase mainly reflected a financial crisis at the end of the 1990s. The fourth panel shows that

Figure 3: **Colombian Aggregates**



Notes: In all panels, 1991 is marked as the reform year. See text for details about variables. Missing data points were unavailable. Data sources for each panel in clockwise order: 1) DANE Annual Survey of Manufacturers. Pre-1991 series are based on own calculations from the micro-data, post-1991 series has been obtained from DANE; 2) Inter-American Development Bank (2004). See footnote 24 for the definition of job turnover; 3) International Monetary Fund (2011); 4) Blue line: urban employment share of Manufacturing (ISIC-Rev.2 code 3), black line: combined employment shares of Wholesale and Retail Trade and Restaurants and Hotels, and Community, Social and Personal Services (ISIC-Rev.2 codes 6 and 9), ILOSTAT Database; 5) Approximate numbers based on Figure 3.1 in Mondragón-Vélez and Pena (2010) , Colombian Household Survey (Encuesta Nacional de Hogares, ENH); 6) World Bank (2013).

after reforms, the manufacturing share of urban employment dropped from 24 percent in 1991 to roughly 20 percent in 2000. The corresponding increase in the service sector was largely driven by two sub-sectors, both of which exhibit a high level of self-employment: wholesale-retail trade and personal services (Mondragón-Vélez and Pena 2010). So the sustained increase in the share of self-employed urban workers (fifth panel) can be taken as a sign of weakening demand in formal labor markets. Finally, the sixth panel shows that over the same time period, the Gini coefficient for Colombia rose from roughly 53 percent to roughly 58 percent.

These aggregate trends were accompanied by a dramatic shift in the plant size distribution. Figure 4 shows the size distribution of manufacturing plants in the 1980s (black bars) and 2000s (white bars). The average size increased from 45 to 60 workers, and the proportion of plants with more than 100 workers increased from 15 percent to 22 percent.

In sum, Colombia experienced a significant shift in its manufacturing plant size distribution and an overall decline in manufacturing employment. Manufacturing jobs also became more unstable as job turnover rates increased. Unemployment, self-employment, and wage inequality also increased. We now investigate how, in the context of our model, these changes might be linked to the changes in tariffs, firing costs and foreign market conditions that Colombian firms experienced during the late 1980s and early 1990s.

3.2 Fitting the model to the data

Model Period Assuming that Colombia was in a steady state prior to reforms, we fit our model to annual data from 1981-1990. In doing so, we treat all plants as single-plant firms. Also, in order to exploit "control function" techniques (discussed below), we match the periodicity of our model to the periodicity of our data. This means imposing that unemployment spells occur in one-year increments, which is longer than some calibrated labor search models for the U.S. economy have presumed (e.g., Elsby and Michaels, 2013; Cooper et al., 2007). Nonetheless, since the average unemployment spell in urban Colombian labor markets is around 11 months (page 16 in Medina et al. 2013), this drawback does not strike us as critical. We will return to the issue of periodicity when we discuss the robustness of our findings.

Parameters not estimated Several parameters are not identified by the model; these we take from external sources. The real borrowing rate in Colombia fluctuated around 15 percent between the late 1980s and early 2000s, so we set $r = 0.15$ (Bond et al., forthcoming). The average share of services in Colombian GDP during the sample period was 0.48, so this is our estimate for γ.[17] Heckman and Pages (2000) estimate that dismissal costs amounted to 6 months' wages in 1990 (their Figure 1), so we fix firings costs at $c_f = 0.6$ in the benchmark economy.[18] Eaton and Kortum (2002) estimate that the tariff equivalent of iceberg costs falls between 123 percent and 174 percent, so we choose our pre-reform value of $\tau_c - 1$ to be 1.50. Finally, we take our estimate of the pre-reform nominal tariff rate, $\tau_a - 1 = 0.21$, from Goldberg and Pavcnik (2004).

The estimator This leaves us with 16 parameters to estimate, collected in the vector

$$\Omega = (\sigma, \alpha, \rho, \sigma_z, \beta, \theta, \delta, \lambda_1, \lambda_2, b, l_e, c_h, c_p, c_x, c_e, D_F^*).$$

These we estimate using the method of simulated moments (Gouriéroux and Monfort 1996). Specifically, let $\bar{m}$ be a vector of sample statistics that our model is designed to explain and define $m(\Omega)$ as the vector of model-based counterparts to these sample statistics. Our estimator is then given by

$$\widehat{\Omega} = \arg\min \left(\bar{m} - m(\Omega)\right)' \hat{W} \left(\bar{m} - m(\Omega)\right),$$

[17] Source: ICP Table 7 downloaded from http://www.eclac.cl/deype/PCI_resultados/eng/index.htm).

[18] In the benchmark economy average wages are about 1.2. Since the model period is a year, 6 months' wages is $c_f = 0.6$. Note that c_f is not a parameter that we can select without running the model as its value depends on average wage in the benchmark economy. Since we match it exactly, however, we do not report its standard errors.

where $\hat{W}$ is a bootstrapped estimate of $[var(\bar{m})]^{-1}$ with off-diagonal elements set to zero.[19]

The sample statistics The vector $\bar{m}$ and the associated weighting matrix are based on plant-level panel data from Colombia. These data are annual observations on all manufacturing plants with at least 10 workers, covering the 1981-1990 period.[20]

Table 2 lists the elements of $\bar{m}$, grouped according to the type of information they convey. The first group consists of means, variances, and covariances for the vector $(\ln l_t, \ln G_t, \mathcal{I}_t^x, \ln l_{t+1}, \ln G_{t+1}, \mathcal{I}_{t+1}^x)$.[21] Gross revenues G are gross sales, expressed in thousands of 1977 pesos. The indicator $\mathcal{I}^x$ takes a value of unity for those plant-year observations with positive exports. Finally, since workers are all identical in the model economy, we control for the effects of worker heterogeneity on output by measuring the labor input l in terms of "effective worker" units (see Appendix 5.2 for details).

The second and third groups of moments in $\bar{m}$ include quintiles of the plant size distribution and the average rate of employment growth among expanding plants within each size category, respectively. We target employment growth among expanding plants since due to linear firing costs, contractions are not gradual in the model. Quintiles are based on effective employment levels, l, and constructed using the pooled panel of plants.[22] Employment growth rates for quintile j are constructed as cross-plant averages of $(l_{t+1} - l_t)/\left[\frac{1}{2}(l_{t+1} + l_t)\right]$, including only expanding plants that were in quintile j at the beginning of the period. New plants are included in these growth rates, and are treated as having an initial employment of zero.

The fourth group in $\bar{m}$ contains aggregate statistics for the pooled sample of plants. These include the job turnover rate, the plant exit rate, and the standard deviation in effective wages. Job turnover is a cross-year average of the annual turnover rate, net of aggregate

[19] Setting off-diagonal terms to zero improves the stability of our estimator while maintaining consistency and keeping it independent of units of measurement. Examples of other studies employing the same strategy include Lee and Wolpin (2006) and Dix-Carneiro (2014).

[20] The data were collected by Colombia's National Statistics Department (DANE) and cleaned as described in Roberts (1996). They cover 88,815 plant-observations during the sample period. Estimates of $\widehat{var}(\bar{m})$ are generated by bootstrapping the sample.

[21] In a stationary equilibrium, $E(\ln l_{t+1}, \ln G_{t+1}, \mathcal{I}_{t+1}^x) = E(\ln l_t, \ln G_t, \mathcal{I}_t^x)$ and $cov(\ln l_{t+1}, \ln G_{t+1}, \mathcal{I}_{t+1}^x) = cov(\ln l_t, \ln G_t, \mathcal{I}_t^x)$. We therefore exclude $E(\ln l_{t+1}, \ln G_{t+1}, \mathcal{I}_{t+1}^x)$ and $cov(\ln l_{t+1}, \ln G_{t+1}, \mathcal{I}_{t+1}^x)$ from our moment vector. We also drop $var(I_t^x)$ because it is redundant: the variance of a Bernoulli random variable depends solely on its mean. This leaves 3 means, 2 variances, and 12 covariances.

[22] While our estimation allows l_e (the size of entering plants) to be arbitrarily small, our database does not cover plants with less than 10 workers. This means that plants appearing in the database for the first time can either be plants crossing the 10-worker threshold from below, or plants in their first year of operation. We apply the same truncation to our simulated moments. This means, for example, that statistics describing the smallest quintile characterize the smallest quintile among *observed* producers.

employment growth or contraction.[23] The plant exit rate is the fraction of plants that exit the panel in year t, averaged over the 10-year sample period. Finally, the standard deviation in effective wages is constructed as the cross-plant standard deviation of the log of real payments to labor (wages and benefits) per effective worker. Given that our measure of effective workers has been adjusted for workforce composition, this measure of wage dispersion controls for observable worker characteristics to the extent possible. Unavoidably, it partly reflects variation in unobservable worker characteristics. But this latter source of noise is averaged across individual workers within a firm, and thus is hopefully relatively modest.[24]

The last two elements of $\bar{m}$ are not simple descriptive statistics. Rather, they are sample-based estimates of $\left(\frac{\sigma-1}{\sigma}\right)(1-\alpha)$ and d_F obtained by applying the logic of Olley and Pakes (1996) to the gross revenue function. By including these statistics in the moment vector rather than treating them as fixed parameters when estimating Ω, we recognize the effects of their sampling error on $\hat{\Omega}$.[25]

Our approach to estimating these two statistics merits further explanation. By (3) and (12), gross revenues before fixed exporting costs can be written as

$$\ln G_{it} = d_H + \mathcal{I}_{it}^x d_F(\eta_0) + \left(\frac{\sigma-1}{\sigma}\right)\left[\ln z_{it} + \alpha \ln l_{it} + (1-\alpha)\ln m_{it}\right]. \tag{30}$$

Also, among firms that adjust their employment levels, the policy function $l' = L(z', l)$ can be inverted to express z' as a monotonic function of l' : $\ln z' = g(\ln l, \ln l')$. This "control function" allows us to eliminate z from (30):

$$\ln G_{it} = \widetilde{d}_H + \mathcal{I}_{it}^x d_F(\eta_0) + \left[\frac{\sigma-1}{\sigma}(1-\alpha)\right]\ln(Pm_{it}) + \varphi(\ln l_{it-1}, \ln l_{it}) + \xi_{it}. \tag{31}$$

Here $\varphi(\ln l_{it-1}, \ln l_{it}) = \frac{\sigma-1}{\sigma}\left[\alpha \ln l_{it} + g(\ln l_{it-1}, \ln l_{it})\right]$ is treated as a flexible function of its arguments, and the intercept $\tilde{d}_H = d_H - (1-\alpha)\frac{\sigma-1}{\sigma}\ln P$ reflects the fact that we have replaced

[23] Let c, e, and d be the set of continuing, entering, and exiting plants, respectively. Also, let i index plants. Our year t job turnover measure is then:

$$X_t = \left(\Sigma_{i\in c}|l_{it} - l_{it-1}| + \Sigma_{i\in e} l_t + \Sigma_{i\in d} l_{t-1} - |\Sigma_i l_{it} - \Sigma_i l_{it-1}|\right)/\Sigma_i l_{it-1},$$

and our turnover statistic is $\frac{1}{10}\sum_{t=1981}^{1990} X_t$. The job turnover numbers in Table 2 are slightly higher than those depicted in Figure 3 for two reasons. First, Figure 3 is based on worker head counts, while our moment is based on effective workers. Second, the turnover rates in Figure 3 are taken from a study limited to establishments with at least 15 workers, while our moment is based on establishments with at least 10 workers. It was not possible to construct Figure 3 using effective workers and a 10 worker cutoff because we did not have access to establishment level data more recent than 1991.

[24] The cross-plant distribution of average wages provides a very natural measure of wage dispersion in a model with homogenous workers. See also Lentz and Mortensen (2008).

[25] The alternative approach, commonly used, is to pre-estimate technology and taste parameters that can be identified without solving the dynamic problem, then treat them as parameters at the computationally intensive stage when parameters identified by the dynamic problem are estimated.

the unobservable m_{it} with observable input expenditures, Pm_{it}. The error term ξ_{it} captures measurement error in $\ln G_{it}$ and any productivity shocks that are unobserved at the time variable inputs and exporting decisions are made. Because ξ_{it} is orthogonal to Pm_{it} and $\mathcal{I}^x_{it}$, we obtain our estimates of $\frac{\sigma-1}{\sigma}(1-\alpha)$ and $d_F(\eta_0)$ by applying least squares to equation (31). Just as Olley and Pakes (1996) excluded observations with zero investment to keep their policy function invertible, we exclude observations for which $l_{it} = l_{it-1}$.[26]

Identification While it is not possible to associate individual parameters in Ω with individual statistics in $\bar{m}$, particular statistics play relatively key roles in identifying particular parameters. We devote this section to a discussion of these relationships.

To begin, the sample-based estimates of $\frac{\sigma-1}{\sigma}(1-\alpha)$ and $d_F(\eta_0)$ provide a basis for inference regarding α and σ. This is because, for any given value of $\left(\frac{\sigma-1}{\sigma}\right)(1-\alpha)$, the elasticity of revenue with respect to labor, $\alpha\Lambda$, increases monotonically in σ. Thus, loosely speaking, the regression of revenue on employment—which is implied by the sample moments $cov(\ln l_t, \ln G_t)$, $var(\ln G_t)$, and $var(\ln l_t)$—pins down σ.[27] And once σ and $\frac{\sigma-1}{\sigma}(1-\alpha)$ are determined, α and Λ are also implied. These moments also discipline θ. As we mentioned above, greater job turnover increases the pool of unemployed workers, and increases the vacancy fill rate, $\phi(V, U)$. Hence θ also helps determine how responsive firms' employment levels are to z.

Next note that by inverting the revenue function, we can express $\ln z_t$ as a function of the data $(\ln l_t, \ln G_t, \mathcal{I}^x_t)$ and several parameters discussed above $(d_F(\eta_0), \alpha, \Lambda)$. Thus, given these parameters, the data vector $(\ln l_t, \ln G_t, \mathcal{I}^x_t, \ln l_{t+1}, \ln G_{t+1}, \mathcal{I}^x_{t+1})$ determines $\ln z_{t+1}$ and $\ln z_t$, and the second moments of this vector imply the parameters of the autoregressive process that generates $\ln z_t$, i.e., ρ and σ_z.

The average level of gross revenues—proxied by $E(\ln G_t)$—helps to identify the fixed cost of operating a firm, c_p. That is, larger fixed costs force low-revenue firms to exit, and thereby increase $E(\ln G_t)$ among survivors. The mean exporting rate $E(\mathcal{I}^x)$ is informative about the fixed costs of exporting, c_x. Also, since the cost of creating a firm, c_e, must match the equilibrium value of entry, the estimated intercept $\widetilde{d}_H$ from (31) helps us to pin down the price level P in the estimation. In turn, this pins down the value of entry $\mathcal{V}_e$ from (21). Further details are provided in Appendix 4.

[26] While a standard application of Olley-Pakes would involve correcting for selection bias, this is not appropriate in the present context. The reason is that our timing assumptions in section 2.7 imply entry and exit decisions are made before the current productivity shock is realized.

[27] In this regression, the error term is a function of z and thus is correlated with labor. But the dependence of l on z is built into our model, so under the maintained hypothesis that the model is correctly specified, there is no simultaneity bias. Put differently, by exploiting our model's structure and assuming constant returns to scale, we avoid the need for a second stage Olley-Pakes step.

The job turnover rate among continuing firms is informative about the general magnitude of hiring costs, which scale with c_h. Similarly, the firm-size-specific job add rates are informative about frictions faced by firms in different states. More precisely, in the absence of labor market frictions, the job turnover rate, the firm size distribution, and the quintile-specific add rates would simply be determined by the productivity process. Deviations from these patterns require adjustment frictions, and quintile-specific patterns require different frictions for firms of different sizes. The parameter λ_1, which governs the convexity of hiring costs, determines the overall level of adjustment frictions. And λ_2, which governs scale economies in hiring costs, determines the relative stability of large versus small firms, as discussed above in Section 2.7. Other things equal, the smaller λ_2 is, the more rapidly the employment growth rate declines with firm size.

Finally, in combination with information on job turnover and hiring rates, the share of employment in the non-traded services sector and the cross-firm dispersion in log wages help to identify the matching function parameter (θ), workers' bargaining power (β), and the value of being unemployed (b). These parameters determine how rents are shared between the workers and the employers in hiring firms, and, as a result, the wage dispersion across firms.

Estimates and model fit Table 1 reports our estimates of Ω, while Table 2 reports the data-based and model-based vectors of statistics, $\bar{m}$ and $m(\widehat{\Omega})$, respectively. Standard errors in Table 1 are constructed using the standard asymptotic variance expression, with $v\hat{a}r(\bar{m})$ bootstrapped from the sample data.[28] Since our data-based moments are calculated from a large survey of plants, sample variation in the moments is small. Almost by construction, this leads to small diagonal elements of $v\hat{a}r(\bar{m})$. Our solution algorithm is summarized in Appendix 4.

Overall, the model fits the data quite well.[29] In particular, it captures the size distribution of firms (Figure 4, first two bars in each bin), the exit rate, the persistence in employment levels, and the variation in growth rates across the plant size distribution. The model underestimates wage dispersion a bit, but this is to be expected, since our data-based measure of wage dispersion controls for only five types of workers, and thus reflects some unobserved worker heterogeneity. In contrast, our model-based dispersion measure is based on the assumption of homogenous effective labor units.

Estimates of b, c_p, c_x and c_e are measured in terms of our numeraire—the price of the service good, or equivalently, the average annual service sector wage. We calculate this

[28] Specifically, the variance covariance matrix is $(J'WJ)^{-1}(J'W)\hat{Q}(WJ)(J'WJ)^{-1}$, where $J = \partial\Omega'/\partial m$, W is the weighting matrix, and $\hat{Q} = \widehat{cov}(\bar{m} - m(\Omega))$.

[29] At fitted values, the average percentage deviation between data- and model-based moments is 12.6 percent.

Table 1: **Parameters Estimated with Simulated Method of Moments**

Parameter	Description	Estimate	Std. Err.
σ	Elasticity of substitution	6.831	0.0404
α	Elasticity of output with respect to labor	0.195	0.0006
ρ	Persistence of the z process	0.961	0.0002
σ_z	Standard deviation of the z process	0.135	0.0003
β	Bargaining power of workers	0.457	0.0007
θ	Elasticity of the matching function	1.875	0.0023
δ	Exogenous exit hazard	0.046	0.0001
c_h	Scalar, vacancy cost function	0.696	0.0006
λ_1	Convexity, vacancy cost function	2.085	0.0042
λ_2	Scale effect, vacancy cost function	0.302	0.0007
b	Value of home production	0.403	0.0010
l_e	Initial size of entering firms	6.581	0.0120
c_p	Fixed cost of operating	10.006	0.0189
c_x	Fixed exporting cost	100.23	0.5528
c_e	Entry cost for new firms	25.646	0.209
D_F^*	Foreign market size	1361.239	154.58

to be $w_s = \$3,461$ in 2012 US dollars during the sample period so, expressed in dollars, the sunk cost of creating a new firm is $25.65 \times \$3,461 = \$88,771$; the annual fixed cost of operating a business amounts to $10 \times \$3,461 = \$34,610$; and the fixed cost of exporting is $100.23 \times \$3,461 = \$346,880$.[30] (The magnitude of the latter figure reflects the large gap in our sample between the average revenues of exporters and non-exporters.) To put these numbers in context, the mean and median annual sales of a Colombian manufacturing firm during the sample period were $\$4,418,360$ and $\$508,970$, respectively.

Several other features of our results on preferences and technology merit comment. First, our estimate of the elasticity of substitution among differentiated industrial goods, $\sigma = 6.83$, is very much in line with the literature.[31] Second, given our estimates of α and σ, the elasticity of value added with respect to labor is $\alpha\Lambda = 0.53$ (refer to equation 14). This figure falls

[30]The data are expressed in thousands of 1977 pesos. In 1977, there were 46.11 pesos per dollar, and based on the US producer price index, a dollar in 1977 was worth 3.116 in 2012 US dollars. We therefore convert the average industrial wage per effective worker into 2012 US dollars as: $\bar{w} \times 3.116/46.11 = \$4,153$. Then using the ratio of service sector wages to average industrial wages, we compute the service sector wage $w_s = 4,153/1.2 = 3,461$.

[31]Estimates of the elasticity of substitution vary widely; our figure falls somewhere in the middle. For example, using establishment data from Slovenia, De Loecker and Warzynski (2012, Tables 2 and 3) estimate mark-ups ranging from 0.13 to 0.28, implying demand elasticities that range from 2.27 to 8.3. Similarly, using firm-level Indian data, De Loecker et al. (2012) estimate a median mark-up of 1.10, implying a demand elasticity of 11, although they find the distribution of mark-ups is spread over a wide range of values. Using trade data, Baier and Bergstrand (2001) estimate a demand elasticity of 6.43, while Broda and Weinstein (2006) get estimates around 12 for their most disaggregated (10 digit HTS) data.

Table 2: **Data-based versus Simulated Statistics**

Moment	Data	Model	Size Distribution	Data	Model
$E(\ln G_t)$	5.442	5.253	20th percentile cutoff	14.617	15.585
$E(\ln l_t)$	3.622	3.636	40th percentile cutoff	24.010	25.773
$E(\mathcal{I}_t^x)$	0.117	0.108	60th percentile cutoff	41.502	41.432
$var(\ln G_t)$	2.807	3.329	80th percentile cutoff	90.108	79.109
$cov(\ln G_t, \ln l_t)$	1.573	1.788	**Firm Growth Rates**		
$var(\ln l_t)$	1.271	1.219	<20th percentile	1.421	1.234
$cov(\ln G_t, \mathcal{I}_t^x)$	0.230	0.251	20th-40th percentile	0.255	0.271
$cov(\ln l_t, \mathcal{I}_t^x)$	0.152	0.160	40th-60th percentile	0.209	0.183
$cov(\ln G_t, \ln G_{t+1})$	2.702	2.196	60th-80th percentile	0.184	0.151
$cov(\ln G_t, \ln l_{t+1})$	1.538	1.556	**Aggregate Turnover/**		
$cov(\ln G_t, \mathcal{I}_{t+1}^x)$	0.225	0.278	**Wage Dispersion**		
$cov(\ln l_t, \ln G_{t+1})$	1.543	1.394	Firm exit rate	0.108	0.120
$cov(\ln l_t, \ln l_{t+1})$	1.214	1.161	Job turnover	0.198	0.240
$cov(\ln l_t, \mathcal{I}_{t+1}^x)$	0.152	0.185	Std. dev. of log wages	0.461	0.426
$cov(\mathcal{I}_t^x, \ln G_{t+1})$	0.220	0.279	**Olley-Pakes Statistics**		
$cov(\mathcal{I}_t^x, \ln l_{t+1})$	0.149	0.201	$(1-\alpha)\left(\frac{\sigma-1}{\sigma}\right)$	0.685	0.687
$cov(\mathcal{I}_t^x, \mathcal{I}_{t+1}^x)$	0.089	0.073	d_F	0.090	0.094

Notes: All data-based statistics are calculated using Colombian plant-level panel data for the pre-liberalization period, 1981-90. These data were collected by the Colombian National Administrative Department of Statistics (DANE) in its Annual Manufacturer Survey (EAM), which covers all establishments with at least 10 workers.

a bit below the range typically estimated for value-added production functions.[32] Third, we find substantial persistence in the z process ($\rho = 0.96$). This relatively high estimate reflects the fact that, unlike most estimates of productivity processes, we treat capital stocks as fixed upon entry and common across firms. This effectively bundles persistence in employment due to capital stocks into the z process. Fourth, we estimate that about half of the exit that occurs is due to adverse productivity shocks, and half is due to factors outside our model ($\delta = 0.046$). Finally, our model allows us to infer the typical size at which firms enter, recognizing that they do not actually appear in the database until they have acquired 10 workers. This entry size amounts to $l_e = 6.58$ workers.

The remaining parameter estimates in Table 1 concern labor markets. These too are plausible. The returns to home production by unemployed workers is 60 percent lower than the secure wage they could have earned if they had committed to work in the service sector.

[32] Direct comparisons with other recent studies are difficult because most control for capital stocks, and most estimate gross production functions rather than value added functions. One well-known study that does estimate a value-added function is Olley and Pakes (1996). Their preferred estimate of the elasticity of value-added with respect to labor is 0.61–a bit higher than our 0.54. Another well-known study, Ackerberg et al. (2006), reports estimates between 0.75 and 1.0. When comparing to studies that estimate gross physical production functions (correcting for price variation), it is perhaps best to focus on the ratio of the labor elasticity (α) to the materials elasticity. In our model this figure is $\alpha/(1-\alpha) = 0.195/(1-0.195) = 0.24$. In other studies it is either $\alpha_{labor}/\alpha_{materials}$ or $\alpha_{labor}/(\alpha_{materials} + \alpha_{capital})$, depending upon whether one treats capital as a material input. Several recent studies of selected industries find the first measure falls around 0.33 while they find the latter falls around 0.25 (e.g., De Loecker 2011; De Loecker et al. 2012).

Table 3: **Model Implications for Statistics not in $\overline{m}$**

	Data	**Model**
Aggregates		
Revenue share of exports	0.090	0.120
Relative market size (COL/RoW)	0.0057	0.0053
Manufacturing share of employment	0.226	0.248
Unemployment rate	0.108	0.124
Exporters versus Non-exporters		
$\overline{\ln l}_{\mathcal{I}^x=1} - \overline{\ln l}_{\mathcal{I}^x=0}$ (*size premium*)	1.402	1.855
$\overline{\ln w}_{\mathcal{I}^x=1} - \overline{\ln w}_{\mathcal{I}^x=0}$ (*wage premium*)	0.420	0.528
Aggregate employment share of exporters	0.360	0.409
Aggregate revenue share of exporters	0.518	0.629
Wage-Size Relationship		
$corr(w, l)$	0.394	0.102
$\ln w = \alpha + \beta_l \ln l + \beta_x I^x + \varepsilon$		
$\ln l$ coefficient (β_l)	0.201	-0.094
	(0.001)	(0.002)
I^x coefficient (β_x)	0.137	0.702
	(0.004)	(0.005)
R^2	0.295	0.158

Notes: Data-based statistics are constructed using the same panel of establishment used for Table 2. Numbers in parentheses are OLS standard errors.

The matching function parameter, $\theta = 1.88$, is close to the value of 2.16 that Coşar (2013) calibrates using aggregate labor market statistics from Brazil, and not far from the value of 1.27 that den Haan et al. (2000) obtain in calibrating their model to the US economy. The bargaining parameter, $\beta = 0.46$, implies rents are shared nearly equally between workers and firms. Finally, the parameters of the vacancy cost function imply both short-run convexities ($\lambda_1 = 2.09$) and modest scale economies ($\lambda_2 = 0.30$).[33]

Drawing on our discussion in section 2.7 above, we infer that per-worker vacancy posting costs rise substantially with firm size, holding the rate of employment growth constant: $\lambda_1(1-\lambda_2) = 1.46 > 1$. Hence large firms in our model will tend to grow relatively slowly in response to productivity shocks. Since shocks at large and small firms have similar properties—that is, $\rho = 0.96$ implies that mean reversion in z is very slow—this feature of the vacancy cost function plays a dominant role in shaping employment dynamics.

Non-targeted statistics and out-of-sample fit Before discussing policy implications of these estimates, we ask how well the model replicates features of the data that we did not use as a basis for identification. To address this question we construct several additional

[33]Our estimate of λ_1 is consistent with the available evidence on hiring cost convexities (e.g., Merz and Yashiv 2007, and Yashiv 2006). We also come close to satisfying the relationship $\lambda_2 = 1 - 1/\lambda_1$ implied by Nilsen et al.'s (2007) specification.

statistics in Table 3.

We start with the aggregates in the first panel. The pre-reform revenue share of exports, plotted in Figure 3, is 9 percent. In our estimation, we targeted the fraction of firms that export, and the revenue increment due to exporting d_F, but did not explicitly target the revenue share of exports. Yet, the model generates a 12 percent share that is quite close to its empirical level. In the model, $D_H/(k^{\sigma} D_F^*)$ measures the size of domestic expenditures on tradable goods relative to total foreign demand for tradables. We estimate this ratio as 0.0053. While it is hard to find an exact empirical counterpart to this statistic, we calculate Colombia's average GDP relative to the sum of its trade partners' GDP over 1981-1990 and find a value of 0.0057. Another relevant statistic is the employment share of manufacturing, which averaged 0.226 in the pre-reform period. Our model predicts an employment share of $L_q = 0.248$ for the industrial sector, which is again close to the data. Finally, the model-generated unemployment rate among those who search for industrial jobs, $L_u/(L_u + L_q)$, is 0.124, which is quite close to the average Colombian urban unemployment rate of 10.8 percent in the pre-reform period. Our model abstracts from labor market frictions in the service sector in order to focus on intra-sectoral effects in the tradable industry. Unemployment in the data is, however, generated by both sectors. Without detailed data on sectorial job finding and separation rates, it is not possible to gauge the level of search frictions within each sector. Nonetheless, data show us that Colombian transition rates between employment and unemployment are very similar in both sectors (see Appendix 5.2 for further details). This implies similar steady state sectoral unemployment rates, and thus suggests that industrial unemployment figures are representative of the economy overall.[34]

As the second panel of Table 3 shows, Colombian exporters are larger (*size premium*) and pay higher wages (*wage premium*) than non-exporters. Also, as a group they account for more than a third of industrial employment and slightly more than half of total revenues. The model generates all of these patterns, although it overstates the gap between exporters and non-exporters. This tendency to overstate exporter premia while matching other moments reflects the fact that in our model, all firms above a threshold output level are exporters (see equation 11). The contrast between exporters and others could be weakened without sacrificing model fit by adding another source of firm heterogeneity—for example random fixed exporting costs. But the workings of the model that we wish to focus upon would be unaffected, so we opt for simplicity here.

This same deterministic relationship between output and exporting status makes it difficult for our model to generate the observed positive association between size and wages,

[34]This is consistent with evidence on sectoral unemployment rates in U.S. According to the BLS, unemployment rate in manufacturing was 9 and 7.3 percents in 2011 and 2012, respectively. Average unemployment rate in service sectors was 8.2 and 7.3 percent for the same years. See http://www.bls.gov/cps/cpsaat26.pdf, accessed on September 26, 2013.

conditional on exporting status. The third panel of Table 3 reports the wage-size relationship. While our model generates a positive unconditional correlation, adding an exporter dummy to the model-based regression of log wages on log employment turns the coefficient on log employment slightly negative.[35] With the exporter dummy absorbing much of the cross-firm rent variation, two remaining forces are at work in our model. On the one hand, holding productivity and exporting status constant, the marginal revenue product of labor falls with employment, putting downward pressure on wages at large firms. On the other hand, holding exporting status constant, productivity shocks tend to induce a positive correlation between firm size and rents, thus wages tend to be higher at large firms. In the model, the marginal revenue product effect dominates. But in the data, the relation between employment and exports is noisier and additional forces are at work, including unobserved worker heterogeneity and perhaps greater monopoly power among larger firms.[36]

Finally, we ask how well our model does in capturing cross-worker (as opposed to cross-firm) residual wage inequality. Since our establishment survey data do not provide information on individual workers, we are unable to construct our own data-based version of this concept. However, we note that the pre-reform average Gini coefficient was around 0.53 (World Bank, 2013, and figure 3), while our model generates a Gini of 0.26. As an alternative statistic, Attanasio et al. (2004) report the 1984-1990 average of *unconditional* standard deviation of log worker wages as 0.80 using the Colombian Household Survey (their Table 2a). The model counterpart is 0.452. Since both data-based measures incorporate observable and unobservable characteristics of workers and firms, and observable worker characteristics typically explain around a third of the wage variation in Mincer regressions (Mortensen 2005), it seems reasonable that our model generates around half of total dispersion by these measures.

4 Simulated Effects of Globalization and Reforms

4.1 The experiments

We are now prepared to examine the effects of reforms and falling trade frictions in our estimated model. Our aim is to determine the extent to which these simulated effects can capture the long-term changes in labor market outcomes documented in Figure 3.[37] Our first

[35] Despite the differences in their wage setting, this result holds equally for both expanding and non-expanding firms.

[36] Similarly, since exporting is strongly correlated with productivity, a negative size-wage relationship obtains when we condition on productivity (see Table A1 in Appendix 5).

[37] Whenever possible, we focus on the post-2000 period because the early 1990s were too close to the reform years to plausibly approximate a new steady state and the late 1990s were characterized by a financial crisis and recession. However, some series such as job turnover and wage inequality are only available up to 2000.

Table 4: **Effects of Reforms and Globalization**

	Baseline	Reforms	Reforms & Globalization
τ_a (ad valorem tariff rate)	1.21	1.11	1.11
c_f (firing cost)	0.6	0.3	0.3
τ_c (iceberg trade cost)	2.5	2.5	2.1
Size Distribution			
20th percentile	16	17	17
40th percentile	25	25	29
60th percentile	39	40	50
80th percentile	78	81	114
Average firm size	46	49	62
Firm Growth Rates			
$<$20th percentile	1.15	1.14	1.20
20th-40th percentile	0.26	0.26	0.34
40th-60th percentile	0.18	0.19	0.22
60th-80th percentile	0.15	0.16	0.20
Aggregates			
% of firms exporting	1	1.298	2.710
Revenue share of exports	1	1.353	2.497
Exit rate	1	0.866	0.957
Job turnover	1	1.027	1.121
Mass of firms	1	0.929	0.705
Unemployment rate in the industrial sector	1	1.055	1.285
Industrial share of employment	1	0.990	0.939
Standard deviation of log wages (firms)	1	0.999	1.074
Standard deviation of log wages (workers)	1	0.982	0.977
Log 90-10 wage ratio (firms)	1	0.998	1.080
Log 90-10 wage ratio (workers)	1	0.988	1.000
Standard deviation of workers' value (J)	1	1.033	1.165
Log 90-10 ratio of workers' value (J)	1	1.017	1.216
Exchange rate (k)	1	0.987	0.685
Real income	1	1.035	1.280

Note: Aggregate statistics in the bottom panel are normalized by their baseline levels.

experiment reduces tariffs τ_a and firing costs c_f to their actual post-reform levels. Specifically, we cut τ_a from 1.21 to 1.11, and c_f from 0.6 to 0.3. Our second experiment goes beyond the first one by also reducing trade costs, τ_c, from its baseline level of 2.5 to 2.1. This 27 percent drop in $\tau_c - 1$, in combination with the reductions in τ_a and c_f, is chosen to match the observed increase in the aggregate revenue share of exports.

The reduction τ_c captures additional forces of globalization during the period under study. These include the increased income and openness of Colombia's trading partners, improvements in global communications, and general reductions in shipping costs (Hummels, 2007). It also captures the integration of rapidly growing emerging markets into the global economy. We view these shocks as originating beyond Colombia's borders, inasmuch as Latin America in general experienced a surge in trade that roughly matched Colombia's.

4.2 Findings

The results of the first experiment (replicating the reforms) are reported in the second column of Table 4. The top and middle panels report the absolute values of the moments while the bottom panel normalizes the baseline outcomes to one. Note that some of the baseline results slightly differ from their estimated values reported in Table 2.[38]

Note that the reforms increase aggregate real income (I/P^{γ}) by 3.5 percent (see Appendix 3 for the definition of aggregate income I). But they also lead to small increases in job turnover and industrial unemployment, while reducing the size of the industrial sector work force and shifting the firm size distribution rightward. So these policy changes improve average incomes at the expense of job security, but both effects are modest.

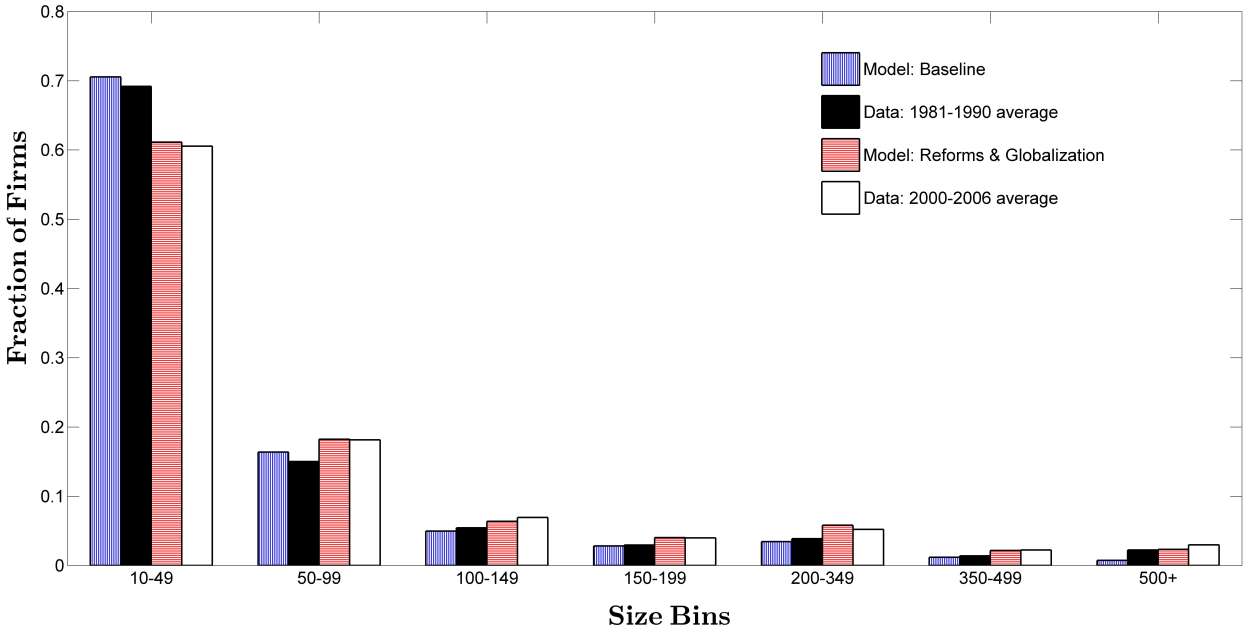
To conserve space we relegate details on the separate effects of τ_a and c_f to Appendix 5. However, we note here that reducing firing costs alone triggers two opposing effects. On the one hand, it induces a 17 percent decline in the exit rate, since these costs create an incentive for firms to shut down. (Recall that, unlike contracting firms, exiting firms are not required to to pay c_f to each displaced worker.) On the other hand, the well-known direct impact of firing cost reductions on job security (e.g., Ljungqvist 2002; Mortensen and Pissarides 1999) leads to an increase in the job turnover rate among continuing producers. These two forces almost cancel each other, so that firing cost reductions alone have very little effect on job turnover, unemployment, or other variables in Table 4. Accordingly, other than the exit rate, the results in the second column are essentially attributable to the reduction in τ_a.

We now turn to our second experiment, which characterizes the combined effects of reforms and globalization. Results are reported in the third column of Table 4. Starting with average firm size, note that our simulation predicts a fairly large increase from 46 to 62 workers as a result of the reductions in τ_a and τ_c. Figure 4 reports the plant size distribution in the data and in the model for pre and post-reform periods. Since the post-reform data are only available in terms of the number of workers, we report the number of workers (not effective workers as we did in Table 2), both for the model and the data (see Appendix 5.2 for details). As Figure 4 shows, our post-reform simulation matches the actual movement in the Colombian plant size distribution quite closely, not only in terms of average size, but also in terms of shape (third and fourth bars in each bin). We match this post-reform rightward shift through the worker reallocation effect emphasized by Melitz (2003).

Turning to job turnover, our simulation predicts an increase of 12 percent, capturing close to half of the 27 percent increase that we observe in Colombia during the post-reform period

[38] This discrepancy is due to computational issues. As we explain in Appendix 4, our quantitative strategy allows us to use an estimation algorithm that is simpler and thus faster than the one we use in simulating the effects of parameter changes. While these two algorithms generate essentially the same results for most moments, some differ due to numerical approximations (such as the first quintile growth rate). The results in Table 4 are generated consistently using the same simulation algorithm.

Figure 4: **Firm Size Distribution: Model vs. Data**



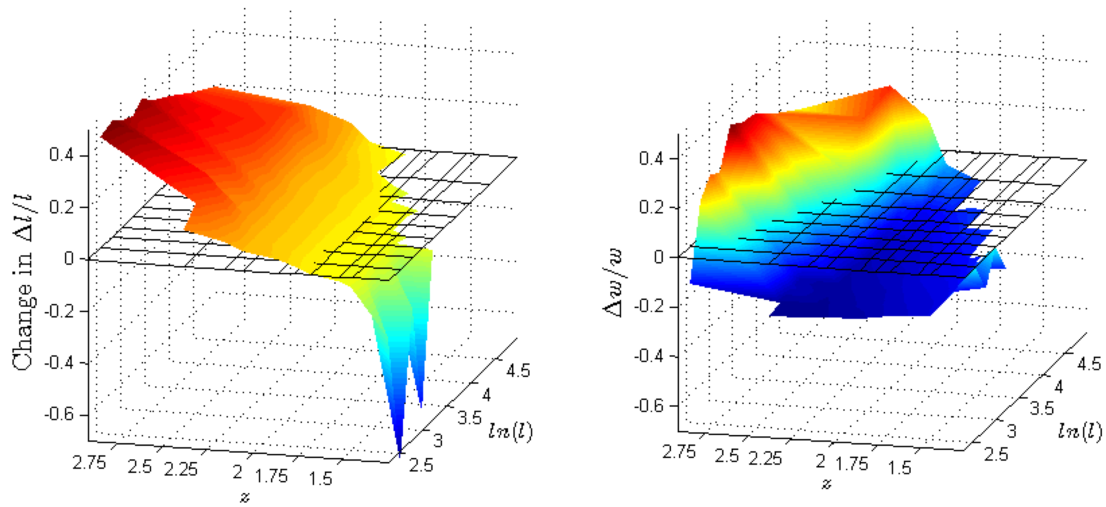
Notes: 1981-1990 average is calculated from plant-level data. 2000-2006 data is obtained from DANE. Since labor in the model is in effective labor units, while the post-reform data is only available in terms of number of employees, we convert model-generated firm size into number of employees using the fit of the two units in the pre-reform plant level data. The details of this procedure are described in Appendix 5.2.

(Panel 2 of Figure 3, and Section 3.1). This reflects the dominance of the sensitivity effect over the distribution effect, as discussed above in section 2.8. That is, without any change in the employment policy function, the rightward shift in the size distribution would have caused a reduction in job turnover, as firms move to a region of the state space with high z values, where the probability of large percentagewise adjustments in l is small. However, firms' employment levels become more sensitive to productivity shocks as reductions in trade costs increase the payoff to hiring among successful firms. This effect is compounded by the reduction in firing costs, which makes contractions less costly. As documented in the third panel ("Firm Growth Rates") in Table 4, firms grow faster in each quintile of the firm size distribution and the increase is particularly pronounced in higher quintiles.

The dominance of the sensitivity effect is graphically depicted in the left panel of Figure 5, which plots the *change* in state-specific employment growth rates ($\Delta l/l$) relative to the pre-reform baseline. With lower tariffs, firing costs, and iceberg trading costs, low z firms shrink much more than they do in the benchmark economy while high z firms experience much larger growth rates (on the order of 20 to 40 percentage points more), especially if they are relatively small.

With more job turnover, contracting firms release more workers to the unemployment

Figure 5: **Employment Policy and Wage Effects**



Notes: z is firm productivity, $\ln(l)$ *is log labor. Both panels display changes from the baseline to the "reforms and globalization" scenario reported in the third column of table 4, plotted as decile averages over the equilibrium productivity and labor distribution of firms in baseline. The left panel plots the percentage change in firm growth rates. The right panel plots the change in wages.*

pool, driving up the number of industrial job seekers. This makes it cheaper for hiring firms to fill vacancies, as the job filling rate ϕ rises from 56 to 68 percent. In turn, the lower marginal cost of hiring makes firms' employment policy functions still more responsive to productivity shocks, completing a feedback loop.

Because of these forces, our model predicts a 28.5 percent increase in the rate of unemployment among workers who participate in the industrial job market. In the data, the unemployment rate rose from an average of 10.8 percent during 1981-1990 to an average of 13 percent during 2000-2006, implying a 20 percent increase. While this is broadly consistent with the increase predicted by our model, we remind the reader that Colombia endured a financial crisis and recession at the end of the 1990s. The effects of this on unemployment probably lingered into the early part of the next decade.

As we argued in Section 3, unemployment alone is an insufficient measure of labor market conditions in developing countries. The declining employment share of manufacturing and the corresponding rise in self-employment, mostly associated with personal services, point to a further deterioration of stable employment opportunities for workers (Panels 4 and 5 of Figure 3). As shown in the third column of Table 4, our model implies a 6 pecent decline in industry's share of employment, thereby accounting for about a third of the observed 17 percent contraction observed in the data (see Section 3.1).[39]

A final outcome of interest is wage inequality, both across firms and across workers. While the average real wage ($\bar{w}/P^{\gamma}$) increases for workers who retain their jobs, differences between the post-reform (second experiment) and pre-reform (baseline) wage schedules depend very much upon employer states (z, l). The right panel of Figure 5 shows changes in firm-level wages from their benchmark values for each (z, l) combination. Wages become more polarized as relatively productive firms benefit from additional export sales and pay higher wages while smaller, less productive firms suffer from increased import competition and lower their wages. This rent polarization is reflected in cross-firm wage dispersion measures (Table 4), and is consistent with the increase in overall inequality observed in Colombia (Panel 6 of Figure 3). Both the standard deviation and the 90th-10th interpercentile differential of firm-level wages (in logs) increase by roughly 8 percent. More dramatically, dispersion in expected lifetime earnings (J) increases 22 percent as workers anticipate they will be paid more variable wages over their life cycles.

Counteracting this cross-firm effect on wage dispersion, the rightward shift in the firm size distribution puts more workers into larger firms. And since wages are equal *within* firms in our model, this tends to reduce inequality across workers. This compositional channel dominates the polarization of wage schedules across firms, reducing the standard deviation of log wages across industrial workers by 2.3 percent and undoing the rise in the interpercentile wage differential. Our model thus predicts little if any effect of increased openness on within-industry residual wage inequality through the rent-sharing mechanism.

Finally, our second experiment (reforms and globalization) predicts sizeable aggregate income gains from globalization through increased selection, market share reallocations, and cheaper intermediates. These effects dominate the upward pressure on our exact price index (P) that results from a fall in the measure of varieties (N), generating a 28 percent increase in real income with respect to the baseline. The net welfare implications of these income gains would ideally be calculated by weighing them against the welfare losses due to greater worker-specific volatility in jobs and wages. But to do so would require introducing risk aversion into the model, which would substantially complicate the analysis.

Robustness As we noted earlier, our estimation strategy dictated that we set the unit of time in our model equal to one year. To explore the implications of this choice, we have calibrated a quarterly version of our model (see Appendix 6 for details). Since this exercise is

[39] Given our simple characterization of service sector production technology, we cannot generate predictions on the share of service sector output coming from own-account (self-employed) producers. In the data, however, around 90 percent of self-employed workers (who are not employers) are in the service sector (Figure 3.4 in Mondragón-Vélez and Peña 2010). Therefore, we find it plausible to associate an expansion of the service sector with increased self-employment.

based on a number of approximations, we view it as only suggestive. Nonetheless, it indicates how our results might have been affected by our choice of periodicity. In particular, other things equal, allowing workers to search more frequently increases their reservation wages. This reduces wage dispersion and tightens the labor market. So, in order to still match the data in terms of wage inequality and job turnover, the quarterly version of our model requires a lower self-employment income (b), a lower matching function elasticity (θ), and a higher elasticity of demand (σ).

These parameter adjustments do not affect the model's qualitative predictions regarding labor market responses to tariffs, iceberg costs and firing costs (τ_a, τ_c, and c_f). Job turnover, firm-level wage dispersion, and industrial sector unemployment all increase in response to reductions in the "reforms and globalization" experiment. Also, as in the annual version of the model, the firm size distribution shifts rightward while there is a sizeable drop in the number of producers. However, the relative magnitudes of some adjustments do depend on the unit of time in the model. Specifically, in the quarterly model, job turnover is less sensitive to globalization, while the increased inequality shows up more through wage dispersion and less through dispersion of worker welfare.

5 Summary

In Latin America and elsewhere, globalization and labor market reforms have been associated with more job turnover, higher unemployment rates, and greater wage inequality. We formulate and estimate a dynamic structural model that links these developments. Our formulation combines ongoing firm-level productivity shocks and Melitz-type (2003) trade effects with labor market search frictions, firing costs, and worker-firm wage bargaining.

Fit to micro data from Colombia, the model delivers several basic messages. First, this country's tariff reductions and labor market reforms in the early 1990s are unlikely to have been the main reason its labor market conditions deteriorated during subsequent decades. Second, reductions in global trade frictions can explain a substantial share of the heightened job turnover, unemployment and wage inequality this country experienced.

Many other countries registered growth rates in merchandise trade similar to Colombia's over the past two decades, even without major commercial policy reforms. To the extent that these surges were mainly caused by the international integration of product markets, globalization may have contributed to similar labor market outcomes in these countries as well.

In principle, our analysis could be extended in several directions. First, incorporating worker heterogeneity would permit us to link openness with wage effects among workers with different skills and/or at different stages in their careers. Second, a more fully-articulated representation of the service sector would allow us to better characterize economywide patterns of unemployment and perhaps also explicitly deal with informal jobs. Finally, introducing risk aversion would permit us to formally link job turnover rates to welfare, and to examine the trade-off between static gains from trade and losses from heightened risks of job loss. We see these extensions as interesting directions for future work.

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On-line Appendices

Appendix 1: The Revenue Function

From (13), the first order condition for firms' optimal m choice is given by

$$Pm = (1-\alpha)\frac{(\sigma-1)}{\sigma}\exp\left[d_H + \mathcal{I}^x d_F(\eta^0)\right]\left(zl^\alpha m^{(1-\alpha)}\right)^{\frac{\sigma-1}{\sigma}},$$

which gives the optimal choice for m as

$$m = \left(\frac{(1-\alpha)}{P}\frac{\sigma-1}{\sigma}\exp\left[d_H + \mathcal{I}^x d_F(\eta^0)\right]\right)^{\frac{\sigma}{\sigma-1}\Lambda}(zl^\alpha)^\Lambda, \tag{32}$$

where $\Lambda = \frac{\sigma-1}{\sigma-(1-\alpha)(\sigma-1)} > 0$. Using this expression to eliminate m from (13), and noting that

$$\frac{\sigma}{\sigma-1}\Lambda = 1 + (1-\alpha)\Lambda,$$

and

$$\frac{\sigma-1}{\sigma} + \Lambda\frac{(1-\alpha)(\sigma-1)}{\sigma} = \frac{(\sigma-1)}{\sigma}\left[1 + (1-\alpha)\Lambda\right] = \Lambda,$$

yields gross revenue at state (z, l):

$$\begin{aligned}
G(z,l) &= \exp\left[d_H + \mathcal{I}^x d_F(\eta^0)\right](zl^\alpha)^{\frac{\sigma-1}{\sigma}}\left\{\left(\frac{(1-\alpha)}{P}\frac{\sigma-1}{\sigma}\exp\left[d_H + \mathcal{I}^x d_F(\eta^0)\right]\right)^{\frac{\sigma}{\sigma-1}\Lambda}(zl^\alpha)^\Lambda\right\}^{\frac{(1-\alpha)(\sigma-1)}{\sigma}}, \\
&= \exp\left[d_H + \mathcal{I}^x d_F(\eta^0)\right](zl^\alpha)^{\frac{\sigma-1}{\sigma}}\left[\left(\frac{1-\alpha}{P}\right)\left(\frac{\sigma-1}{\sigma}\right)\exp\left(d_H + \mathcal{I}^x d_F(\eta^0)\right)\right]^{(1-\alpha)\Lambda}(zl^\alpha)^{\Lambda\frac{(1-\alpha)(\sigma-1)}{\sigma}}, \\
&= P^{-(1-\alpha)\Lambda}\left[(1-\alpha)\left(\frac{\sigma-1}{\sigma}\right)\right]^{(1-\alpha)\Lambda}\left(\exp\left[d_H + \mathcal{I}^x d_F(\eta^0)\right]\right)^{\frac{\sigma}{\sigma-1}\Lambda}(zl^\alpha)^\Lambda.
\end{aligned}$$

We can now derive a parameterized version of the net revenue function (14). From (32), optimal expenditures on intermediate inputs are:

$$Pm = P^{-(1-\alpha)\Lambda}\left[\left((1-\alpha)\frac{\sigma-1}{\sigma}\right)\exp\left[d_H + \mathcal{I}^x d_F(\eta^0)\right]\right]^{\frac{\sigma}{\sigma-1}\Lambda}(zl^\alpha)^\Lambda.$$

Subtracting this expression and fixed exporting costs from gross revenues yields:

$$
\begin{aligned}
R(z,l) &= G(z,l) - Pm - c_x \mathcal{I}^x \\
&= \left[1-(1-\alpha)\frac{\sigma-1}{\sigma}\right] P^{-(1-\alpha)\Lambda}\left((1-\alpha)\frac{\sigma-1}{\sigma}\right)^{(1-\alpha)\Lambda}\left(\exp\left[d_H+\mathcal{I}^x d_F(\eta^0)\right]\right)^{\frac{\sigma}{\sigma-1}\Lambda}(zl^\alpha)^\Lambda - c_x\mathcal{I}^x \\
&= \left[\frac{\sigma-(1-\alpha)(\sigma-1)}{\sigma}\right] P^{-(1-\alpha)\Lambda}\left((1-\alpha)\frac{\sigma-1}{\sigma}\right)^{(1-\alpha)\Lambda}\left(\exp\left[d_H+\mathcal{I}^x d_F(\eta^0)\right]\right)^{\frac{\sigma}{\sigma-1}\Lambda}(zl^\alpha)^\Lambda - c_x\mathcal{I}^x \\
&= P^{-(1-\alpha)\Lambda}\left(\frac{\sigma-1}{\sigma\Lambda}\right)\left((1-\alpha)\frac{\sigma-1}{\sigma}\right)^{(1-\alpha)\Lambda}\left(\exp\left[d_H+\mathcal{I}^x d_F(\eta^0)\right]\right)^{\frac{\sigma}{\sigma-1}\Lambda}(zl^\alpha)^\Lambda - c_x\mathcal{I}^x \\
&= \Theta P^{-(1-\alpha)\Lambda}\exp\left[d_H+\mathcal{I}^x d_F(\eta^0)\right]^{\frac{\sigma}{\sigma-1}\Lambda}(zl^\alpha)^\Lambda - c_x\mathcal{I}^x,
\end{aligned}
$$

where $\Theta = \left(\frac{1}{(1-\alpha)\Lambda}\right)\left[\frac{(1-\alpha)(\sigma-1)}{\sigma}\right]^{\frac{\sigma}{\sigma-1}\Lambda}$.

Appendix 2: The Wage Functions

Hiring Wages In order to characterize wages in hiring firms, we first determine the total surplus for a firm and a worker that are matched in the end-of-period state (z', l'). At the time of bargaining, the surplus that the marginal worker generates for the firm is given by

$$\Pi^{firm}(z', l, l') = \frac{1}{1+r}\left[\frac{\partial \pi(z', l, l')}{\partial l'} + \frac{\partial \mathcal{V}(z', l')}{\partial l'}\right].$$

Note that at the time of bargaining, the vacancy posting and matching process are over and the costs of vacancy postings are sunk. As a result, if bargaining fails, the firm is simply left with fewer workers. Thus we only use the relevant part of the profit function for hiring firms, i.e., when $l' > l$ in (18), denoted by $\pi(z', l, l')$. The surplus that a marginal worker generates consists of two parts: the current increase in the firm's profits, i.e., marginal revenue product net of wages, and the increment to the value of being in state (z', l') at the start of the next period. If the firm does not exit next period, i.e., if $\mathcal{V}(z', l') > 0$, the marginal worker will have a positive value only if the firm expands. Otherwise, the firm will incur the dismissal cost, c_f. If the firm exits, its expected marginal value from the current marginal hire will be zero.

Similarly, the surplus for the marginal worker who is matched by a hiring firm in the end-of-period state (z', l') is

$$\Pi^{work}(z', l') = \frac{1}{1+r}\left[w_h(z', l') + J^e(z', l') - (b + J^o)\right],$$

where the worker enjoys $w_h(z', l')$ in the current period, and starts the next period in a firm with the beginning-of-period state (z', l'). If bargaining fails, the worker remains unemployed this period, engages in home production of b, and starts the next period in state o.

The worker and firm split the total surplus by Nash bargaining where the bargaining power of the firm is given by β:

$$\beta\Pi^{firm}(z', l, l') = (1-\beta)\Pi^{wor\ker}(z', l').$$

Wages are thus determined as a solution to the following equation:

$$\beta \left[\frac{\partial \pi(z', l, l')}{\partial l'} + \frac{\partial \mathcal{V}(z', l')}{\partial l'} \right] = (1 - \beta) \left[w_h(z', l') + J^e(z', l') - (b + J^o) \right]. \tag{33}$$

Note that theoretically we cannot rule out the case in which a firm hires in the current period and exits at the beginning of the next period. The bargaining outcome depends on the decision to exit or continue which is made by the time of bargaining. We analyze these two cases separately.

1. Exiting firms: If the firm is going to exit next period, i.e., $\mathcal{I}^c(z', l') = 0$, we have $\partial \mathcal{V}(z', l')/\partial l' = 0$ and $J^e(z', l') = J^u$ from the definition of J^e. In this case, $\partial \mathcal{V}(z', l')/\partial l'$ cancels with $J^e - J^o$ in (33) since $J^o = J^u$ in equilibrium. We are left with

   $$\beta \frac{\partial \pi(z', l, l')}{\partial l'} = (1 - \beta)[w_h(z', l') - b]. \tag{34}$$

   Using the definition of $\pi(z', l')$ from (18), and rearranging terms, equation (34) becomes

   $$\frac{\partial w_h(z', l')}{\partial l'} \beta l' + w_h(z', l') - \beta \frac{\partial R(z', l')}{\partial l'} - (1 - \beta) b = 0,$$

   which is the same as equation (10) in Bertola and Garibaldi (2001). From (14) we have:

   $$\frac{\partial R(z', l')}{\partial l'} = \Delta \alpha \Lambda \left(z'\right)^{\Lambda} \left(l'\right)^{\alpha \Lambda - 1}.$$

   Here, we suppressed the dependence of $\Delta(\cdot)$ on l' since $\partial \Delta / \partial l' = 0$ if the firm's exporting decision does not depend on the marginal worker. Since workers bargain individually and simultaneously with the firm, no single worker will be taken as the marginal worker for the export decision. Accordingly, retracing Bertola and Garibaldi's (2001) derivation we obtain:

   $$\begin{aligned} w_h(z', l') &= (1 - \beta) b + l^{-\frac{1}{\beta}} \int_0^l u^{\frac{1-\beta}{\beta}} \Delta \alpha \Lambda \left(z'\right)^{\Lambda} \left(l'\right)^{\alpha \Lambda - 1} du \\ &= (1 - \beta) b + \Delta \alpha \Lambda \left(z'\right)^{\Lambda} \left(l'\right)^{-\frac{1}{\beta}} \int_0^{l'} u^{\frac{1}{\beta} + \alpha \Lambda - 2} du \\ &= (1 - \beta) b + \frac{1}{\frac{1}{\beta} + \alpha \Lambda - 1} \Delta \alpha \Lambda \left(z'\right)^{\Lambda} \left(l'\right)^{\alpha \Lambda - 1} \\ &= (1 - \beta) b + \frac{\beta}{1 - \beta + \alpha \beta \Lambda} \underbrace{\Delta \alpha \Lambda \left(z'\right)^{\Lambda} \left(l'\right)^{\alpha \Lambda - 1}}_{= \partial R(z', l') / \partial l'}. \end{aligned}$$

   In this case, the worker is paid a fraction of her marginal revenue plus her share of the outside option b.

2. Continuing Firms: In this case, we have $\mathcal{V}(z', l') > 0$. There is an expected gain from keeping the marginal worker because of the possibility of further hiring next period. The worker's expected gain in the beginning of the next period (when she still has

a chance to leave the firm and search) is $J^e(z', l') - J^u$. In line with Bertola and Caballero (1994) and Bertola and Garibaldi (2001), we assume that the firm-worker pair shares the expected match surplus in the same manner they split current rents, i.e., $J^e(z', l') - J^u$ cancels with the expected gain of the firm in (33). In the event of a contraction, however, the firm cannot enforce contracts that require laid-off workers to pay their share of firing costs. As a result, expected firing costs, $P_f(z', l')c_f$, are subtracted from firm surplus in the current period:

$$\beta \left[\frac{\partial \pi(z', l, l')}{\partial l'} - P_f(z', l')c_f \right] = (1 - \beta)[w_h(z', l') - b],$$

Conditional on the firm not hiring, the possibility of losing one's job, $p_f(z', l)$, is

$$p_f(z', l) = \frac{l - L(z', l)}{l},$$

and the probability of being fired next period is then given by

$$P_f(z', l') = E_{z''/z'} \left\{ \left[1 - \mathcal{I}^h(z'', l') \right] p_f(z'', l') \right\}.$$

The wage schedule for expanding firms that will stay in the market next period is then given by

$$w_h(z', l') = (1 - \beta)b + \frac{\beta}{1 - \beta + \alpha\beta\Lambda} \Delta\alpha\Lambda \left(z'\right)^{\Lambda} \left(l'\right)^{\alpha\Lambda - 1} - \beta P_f(z', l')c_f.$$

Firing Wages To derive the firing wage schedule, we begin by writing the value of employment at a firing firm in the interim stage as

$$J_f^e(z', l) = \frac{1}{1 + r} \left[p_f(z', l)(1 + r)J^u + (1 - p_f(z', l)) \left(w_f(z', l') + J^e(z', l') \right) \right],$$

where $l' = L(z', l)$. This expression reflects the fact that workers who are not fired are paid just enough to retain them. Since workers are indifferent between staying and leaving, the two outcomes inside the bracket have equal value, i.e.,

$$w_f(z', l') + J^e(z', l') = (1 + r)J^u,$$

which yields the wage schedule according to which workers in firing firms are paid:

$$w_f(z', l') = rJ^u - [J^e(z', l') - J^u].$$

Appendix 3: Steady State Equilibrium

Let the transition density of the Markov process on z be denoted by $h(z'|z)$. Given a measure of aggregate expenditure abroad denominated in foreign currency, D_F^*, a steady state equilibrium for a small open economy consists of: a measure of domestic differentiated goods N_H; an exact price index for the composite good P; an aggregate domestic demand index for industrial goods D_H; aggregate income I; a measure of workforce in services L_s; a measure workers in differentiated goods sector L_q; a measure of workers searching for jobs in the industrial sector U; a measure of unemployed workers L_u; the job finding rate $\widetilde{\phi}$; the vacancy filling rate ϕ; the exit rate μ_{exit}; the fraction of firms exporting μ_x; the measure of entrants M; the value and associated policy functions $\mathcal{V}(z,l)$, $L(z,l)$, $\mathcal{I}^h(z,l)$, $\mathcal{I}^c(z,l)$, $\mathcal{I}^x(z,l)$, J^o, J^u, J^s, and J^e; the wage schedules $w_h(z,l)$ and $w_f(z,l)$; the exchange rate k; and end-of period and interim distributions $\psi(z,l)$ and $\widetilde{\psi}(z,l)$ such that:

1. **Steady state distributions:** In equilibrium, $\psi(z,l)$ and $\widetilde{\psi}(z',l)$ reproduce themselves through the Markov processes on z, the policy functions, and the productivity draws upon entry. In order to define the interim distribution, $\widetilde{\psi}(z,l)$, let $\widetilde{\widetilde{\psi}}(z',l)$ be the interim frequency measure of firms defined as

$$\widetilde{\widetilde{\psi}}(z',l) = \begin{cases} \int_z h(z'|z)\psi(z,l)\mathcal{I}^c(z,l)dz \text{ if } l \neq l_e \\ \psi_e(z') + \int_z h(z'|z)\psi(z,l)\mathcal{I}^c(z,l)dz \text{ if } l = l_e \end{cases}.$$

Then, $\widetilde{\psi}(z',l)$ is given by

$$\widetilde{\psi}(z',l) = \frac{\widetilde{\widetilde{\psi}}(z',l)}{\int_{z'}\int_l \widetilde{\widetilde{\psi}}(z',l)dz'dl},$$

while the end-of period distribution is

$$\psi(z',l') = \frac{\int_l \widetilde{\psi}(z',l)\mathcal{I}_{(L(z',l),l')}dl}{\int_{z'}\int_l \widetilde{\psi}(z',l)\mathcal{I}_{(L(z',l),l')}dz'dl},$$

where $\mathcal{I}_{(L(z',l),l')}$ is an indicator function with $\mathcal{I}_{(L(z',l),l')} = 1$ if $L(z',l) = l'$.

2. **Market clearance in the service sector:** Demand for services comes from two sources: consumers spend a $(1-\gamma)$ fraction of aggregate income I on it, and firms demand it to pay their fixed operation and exporting costs, as well as labor adjustment and market entry costs. Aggregate income I itself is the sum of wage income earned by service and industrial sector workers, market services supplied by unemployed workers, tariff revenues rebated to worker-consumers, and aggregate profits in the industrial sector distributed to worker-consumers who own the firms.

The average labor adjustment cost is given by

$$\overline{c} = \int_z \int_l C(l, L(z,l))\widetilde{\psi}(z,l)dldz.$$

The market clearance condition is then given by

$$L_s + bL_u = (1-\gamma)I + N_H(\overline{c} + c_p + \mu_x c_x) + Mc_e.$$

3. **Labor market clearing:** Total production employment in the industrial sector is

given by

$$L_q = N_H \bar{l} = N_H \int_z \int_l l\psi(z,l) dldz,$$

where

$$\bar{l} = \int_z \int_l l\psi(z,l) dldz \tag{35}$$

is the sector's average employment. Every period a fraction μ_l of workers in that sector is laid off due to exits and downsizing:

$$\mu_l = \frac{\int_z \int_l [1 - \mathcal{I}^c(z,l)] l\psi(z,l) dldz + \int_z \int_l \mathcal{I}^c(z,l)\mathcal{I}^f(z,l)[l - L(z,l)]\psi(z,l) dldz}{\int_z \int_l l\psi(z,l) dldz}$$

Then, the equilibrium flow condition is

$$U\widetilde{\phi} = L_q \mu_l.$$

In equilibrium, a measure of $L_u = (1 - \widetilde{\phi})U$ of workers who search do not find a job, and labor market clearing condition is given by

$$1 = L_s + L_q + L_u.$$

On the vacancies side, the aggregate number of vacancies in this economy is given by

$$V = N_H \int_z \int_l v(z,l)\mathcal{I}^h(z,l) \frac{\widetilde{\psi}(z,l)}{\mu_h} dldz = N_H \overline{v},$$

where

$$\overline{v} = N_H \int_z \int_l v(z,l)\mathcal{I}^h(z,l) \frac{\widetilde{\psi}(z,l)}{\mu_h} dldz, \tag{36}$$

is the average level of vacancies, and μ_h is the fraction of hiring firms:

$$\mu_h = \int_z \int_l \mathcal{I}^h(z,l)\widetilde{\psi}(z,l) dldz.$$

The total number of vacancies, V, together with U, determines matching probabilities $\phi(V,U)$ and $\widetilde{\phi}(V,U)$ that firms and workers take as given.

4. **Firm turnover:** In equilibrium, there is a positive mass of entry M every period so that the free entry condition (21) holds with equality. The fraction of firms exiting is implied by the steady state distribution and the exit policy function,

$$\mu_{exit} = \int_z \int_l [1 - \mathcal{I}^c(z,l)]\psi(z,l) dldz + \delta,$$

and measure of exits equals that of entrants,

$$M = \mu_{exit} N_H.$$

5. **Trade balance:** Adding up final and intermediate demand, total domestic expenditures on imported varieties equals $D_H (\tau_a \tau_c k)^{1-\sigma}$. Taking the import tariff into account, domestic demand for foreign currency (expressed in domestic currency) is thus $\frac{D_H(\tau_a \tau_c k)^{1-\sigma}}{\tau_a} = D_H \tau_a^{-\sigma} (\tau_c k)^{1-\sigma}$. Tariff revenue is given by $D_H \tau_a^{-\sigma} (\tau_c k)^{1-\sigma} (\tau_a - 1)$, and is returned to worker-consumers in the form of lump-sum transfers. Total export revenues are $\frac{k D_F^* P_X^{*1-\sigma}}{\tau_c}$ with the foreign market price index for exported goods P_X^* as defined in Section 2.3. Trade is balance given by

$$\underbrace{\frac{D_H (\tau_a \tau_c k)^{1-\sigma}}{\tau_a}}_{\text{domestic demand for foreign currency}} = \underbrace{\frac{k D_F^* P_X^{*1-\sigma}}{\tau_c}}_{\text{export revenue}}.$$

   The exchange rate k moves to ensure that this condition holds. Balanced trade ensures that national income matches national expenditure.

6. Workers are indifferent between taking a certain job in the undifferentiated sector and searching for a job in the industrial sector: $J^o = J^s = J^u$.

Appendix 4: Numerical Solution Algorithm

To compute the value functions, we discretize the state space on a log scale using 550 grid points for employment and 60 grid points for productivity. We set the maximum firm size as 2000 workers and numerically check that this is not restrictive. In the steady state, a negligible fraction of firms reaches this size, which is also the case in the data. The algorithm works as follows:

1. Formulate guesses for $D_H, w_f(z,l), w_h(z,l), d_F$ and ϕ. Given ϕ, calculate $\widetilde{\phi} = (1-\phi^\theta)^{1/\theta}$.

2. Given D_H, $w_f(z,l), d_F$, ϕ and $w_h(z,l)$, calculate the value function for the firm, $\mathcal{V}(z,l)$, using equation (19) and find the associated decision rules for exiting, hiring, and exporting. Calculate the expected value of entry, $\mathcal{V}_e$, using equation (21). Compare $\mathcal{V}_e$ with c_e. If $\mathcal{V}_e > c_e$, decrease D_H (to make entry less valuable) and if $\mathcal{V}_e < c_e$, increase D_H (to make entry more valuable). Go back to Step 1 with the updated value of D_H and repeat until D_H converges.

3. Given $w_f(z,l), d_F$, ϕ and the converged value of D_H from Step 2, update $w_f(z,l)$. To do this, first calculate $J^e(z',l')$ using equations (24) and (27), and imposing the equilibrium condition $J^u = J^o$. Given $J^e(z,l)$, update firing wage schedule using equation (29). Compare the updated firing wage schedule with the initial guess. If they are not close enough go back to Step 1 with the new firing wage schedule and repeat Steps 1 to 3 until w_f converges. Note that if firing wages are too high, then $J^e(z,l)$—the value of being in a firm at the start of a period— is high, since the firm is less likely to fire workers. A high value of $J^e(z,l)$, however, lowers firing wages. Similarly, if firing wages are too low, then J^e is low, which pushes firing wages up.

4. Given d_F and ϕ, the converged value of D_H from step 2, and the converged value of $w_f(z,l)$ from Step 3, update $w_h(z,l)$ using equation equation (28).

5. Given ϕ, the converged value of D_H from Step 2, the converged value of $w_f(z,l)$ from Step 3, and the converged value of $w_h(z,l)$ from step 4, calculate the trade balance. To do this:

   (a) Given firms' decisions, calculate $\psi(z,l)$ and $\widetilde{\psi}(z,l)$, the stationary probability distributions over (z,l) at the end and interim states, respectively.

   (b) Given $\widetilde{\psi}(z,l)$, calculate the average number of vacancies and the average employment in the industrial sector using equations (35) and (36).

   (c) Take a guess for N_H. Given N_H and $\overline{v}$, calculate the mass of unemployed U in the industrial sector from

   $$\phi(V,U) = \frac{M(V,U)}{V} = \frac{U}{((vN_H)^\theta + U^\theta)^{1/\theta}},$$

   which is one equation in one unknown. Given U, calculate $L_u = (1-\widetilde{\phi})U$. Then, given $\overline{l}$, the size employment in the service sector is given by $L_s = 1 - L_u - N_H\overline{l}$. Given N_H, L_s, L_u, M (mass of entrants), and I (aggregate income), check if supply and demand are equal in the service sector:

   $$\underbrace{L_s + bL_u}_{\text{supply}} = \underbrace{(1-\gamma)\, I + N_H(\overline{c} + c_p + \mu_x c_x) + Mc_e}_{\text{demand}}.$$

   Update N_H until supply equals demand.

   (d) Given the value of N_H from Step 4c, calculate exports and imports. If exports are larger than imports, lower d_F; if exports are less than imports, increase d_F. Go back to Step 1 with the updated value of d_F, and repeat until convergence.

6. Given the converged value of D_H from Step 2, the converged value of $w_f(z,l)$ from Step 3, the converged value of $w_h(z,l)$ from Step 4, and the converged value of d_F from Step 5, update ϕ. In order to do that, first calculate EJ_h^e using (24). Given EJ_h^e and $\widetilde{\phi}$, calculate J^u using (23). If $J^o > J^u$, increase ϕ (to attract workers to the differentiated goods sector) and if $J^o < J^u$, we lower ϕ (to make the differentiated goods sector less attractive). Go back to Step 2, and repeat until ϕ converges.

Estimation Procedure In our policy experiments, we use the complete algorithm above to compute equilibrium outcomes for given a set of parameters, including the cost of entry c_e. In these experiments, both d_F and D_H are equilibrium objects that respond to changes in τ_a, τ_c and c_f. While estimating the model, however, we use the Olley-Pakes intercept $\tilde{d}_H$ estimated from (31) to calculate firms' net revenue schedule $R(\cdot)$. Similarly, we treat d_F as a moment to be matched: given $\tilde{d}_H$ and the simulated value of the foreign market size parameter D_F^*, we calculate η using equation (10), which allows to use the implied d_F directly in our solution algorithm. The equilibrium price level P and exchange rate k can easily be solved in equilibrium so that trade balance holds and $\tilde{d}_H$ is consistent with D_H. Also, assuming that the economy is in a steady state with positive entry, we back out c_e by setting it equal to the equilibrium value of entry $\mathcal{V}_e$. This approach to discipline the cost of entry c_e is in line with the quantitative literature (Hopenhayn and Rogerson 1993). These shortcuts allow us to skip Steps 2 and $5d$ in the estimation and considerably reduce the computation time.

Appendix 5: Further Results and Data Sources

5.1 Size-Wage Relation

Table A1 shows the effect of size (measured by the number of workers) and productivity on wages in the data and the model.

Table A1: **Wage-Size Relationship Controlling for Productivity**

$\ln w = \alpha + \beta_l \ln l + \beta_z \ln z + \varepsilon$	**Non-expanding firms**		**Expanding firms**	
	Data	**Model**	**Data**	**Model**
β_l	0.09	-1.89	0.06	-1.3
	(0.002)	(0.048)	(0.002)	(0.018)
β_z	0.541	0.964	0.61	0.465
	(0.005)	(0.117)	(0.005)	(0.071)
R^2	0.5	0.65	0.51	0.58

Notes: l is effective workers defined in Appendix 5.3, w is average wage per effective worker, and z is firm-level productivity implied by equation (30) *given the data and estimated parameter values.*

5.2 Sectoral Labor Flows in Colombia

The Colombian Statistical Agency DANE publishes monthly labor market indicators. We accessed the following link on September 26, 2013:
http://www.dane.gov.co/files/investigaciones/empleo/ech/totalNacional/Mensual/IML_MensualTnacional_01_08.xls.
The file is in Spanish but variable names can be easily translated using online translators. In this file, the worksheets titled "*ocup ramas trim tnal*" indicates monthly sectoral urban employment levels (*Población ocupada según posición ocupacional, CABECERAS*). The worksheet titled "*cesantes ramas trim tnal*" reports last sector of employment for the unemployed (*Población desocupada censate según ramas de actividad anterior, CABECERAS*). We exclude agriculture and mining, and aggregate service industries. The ratio of outflows from employment to unemployment gives sectoral transition rates. For the 2000-2006, average transition rates are 0.137 for manufacturing and 0.148 for services.

5.3 Labor Units

Since workers are all identical in the model economy, we measure the labor input l in terms of "effective worker" units in our estimation. This allows us to control for the effects of worker heterogeneity on output. In the plant-level data we use to estimate our model for the pre-reform period, we observe five categories of workers: managerial, technical, skilled, unskilled, and apprentice. For a given plant-year, effective labor l is the sum of all workers in the plant, each weighted by the average wage (including fringe benefits) for workers in

Table A2: **Isolated Effects**

	Baseline	(I)	(II)	(III)
τ_a (ad valorem tariff rate)	1.21	1.11	1.21	1.21
c_f (firing cost)	0.6	0.6	0.3	0.6
τ_c (iceberg trade cost)	2.5	2.5	2.5	2.1
Size Distribution				
20th percentile	16	17	16	17
40th percentile	25	26	24	27
60th percentile	39	41	38	46
80th percentile	78	84	78	104
Average firm size	46	49	46	57
Firm Growth Rates				
<20th percentile	1.15	1.14	1.12	1.17
20th-40th percentile	0.26	0.28	0.27	0.29
40th-60th percentile	0.18	0.19	0.18	0.22
60th-80th percentile	0.15	0.16	0.16	0.19
Aggregates				
% of firms exporting	1	1.339	0.989	2.191
Revenue share of exports	1	1.339	0.999	2.060
Exit rate	1	0.949	0.832	1.025
Job turnover	1	1.032	1.006	1.096
Mass of firms	1	0.918	1.001	0.764
Unemployment rate in the industrial sector	1	1.076	1.001	1.213
Industrial share of employment	1	0.985	1.002	0.949
Standard deviation of log wages (firms)	1	1.002	0.979	1.035
Standard deviation of log wages (workers)	1	1.002	0.978	0.989
Log 90-10 wage ratio (firms)	1	1.010	0.978	1.045
Log 90-10 wage ratio (workers)	1	1.020	0.981	1.009
Standard deviation of workers' value (J)	1	1.057	1	1.084
Log 90-10 ratio of workers' value (J)	1	1.036	0.973	1.066
Exchange rate (k)	1	0.97	1.05	0.727
Real income	1	1.042	0.993	1.180

Notes: Each column presents the outcomes from an isolated counterfactual scenario. Columns (I): reducing tariffs, Columns (II): reducing firing costs, Columns (III): reducing iceberg trade costs.

its category. For each category of worker, the average wage is based on the mean real wage in the entire 10-year panel and expressed as a ratio to the average real wage for unskilled workers during the same period. Thus wage weights are constant across plants and time, and the only source of variation in l is variation in the employment level of at least one category of worker.

After fitting the model to the pre-reform data, we simulate Colombian reforms and decreased trade costs in Section 4. To evaluate the success of the model in explaining post-reform outcomes, we wish to compare the firm size distribution as predicted by the model to its empirical counterpart. Since we do not have access to the plant-level data from the post-reform period, we don't observe the above-described variables used to construct effective labor. While The Colombian Statistical Agency DANE publishes summary statistics on the size distribution of plants for the 2000-2006 period (`http://www.dane.gov.co/index.`

`php/industria/encuesta-anual-manufacturera-eam`), these are based on the number of total employees.

The following procedure facilitates the comparison of model based and empirical size distributions in both periods (Figure 4). Using the pre-reform plant level, we first fit total number of workers to a polynomial of effective labor l. We then use the coefficients from this regression to convert model-generated effective labor l units to worker count for both the estimated pre-reform and simulated post-reform periods. The blue and red bars in Figure 4—representing the model-based size distributions—are generated using this transformation. The black and white bars representing the empirical size distributions are generated directly from the data using total number of employees.

Appendix 6: Robustness to the Choice of Model Period

To isolate the role of periodicity in driving our results, we hold the estimation strategy fixed by using our estimated revenue function and productivity process to approximate their quarterly counterparts.[40] Then we re-estimated remaining parameters using the same moment vector as in the annual baseline, aggregating simulated quarterly outcomes on flow variables to their annual equivalents, and taking simulated fourth quarter realizations on stock variables to be representative of their annual counterparts (as is done in the annual manufacturing surveys).

Specifically, we kept our estimate of the elasticity of value added with respect to labor ($\alpha\Lambda$) based on annual data, and we chose the root of the quarterly productivity process to replicate our estimate of persistence in the annual process: $\rho_q = \rho_a^{1/4}$. Likewise, we adjusted the discount rate to $r_q = (1 + r_a)^{1/4} - 1$, and we shifted the log revenue function intercept $\widetilde{d}_H$ to put revenue flows on a quarterly basis. Finally, since we saw no good way to approximate the relationship between the variance of the innovations in the annual data ($\sigma^2_{z,a}$) and the variance of the innovations in the quarterly data ($\sigma^2_{z,q}$), we included $\sigma^2_{z,q}$ in the set of parameters to be estimated.

Tables A3 and A4 present the resulting parameter estimates and the fit of the model. The quarterly version doesn't fit as well as the annual baseline, perhaps because of the way we have constrained our revenue function estimates. Nonetheless, the quarterly results do give us some insight into the effects of periodicity choice on parameter estimates and model performance.

The major differences in parameter estimates are in the elasticity of substitution σ, the elasticity of the matching function θ, and the value of home production b. The change in b can be explained by the effect of model frequency on wage inequality. Allowing workers to search more frequently increases their reservation wages, which in turn affects the entire wage schedule. Other things equal, this would lower wage dispersion in the model. So, in order to still match the dispersion of *log* wages, the quarterly calibration lowers the constant term $(1-\beta)b$ in the hiring wage schedule (28). It does so by reducing b from 0.403 to 0.28.[41]

[40] We emphasize "approximate" here because there is no analytical relationship linking the parameters of the annual objects to their quarterly counterparts. The reason is that our revenue function characterizes logs of flows, and thus annual variables are not linear combinations of quarterly variables.

[41] Note that the unit of account is the service sector wage per period, so $b = 0.28$ from the quarterly estimation is directly comparable to the b from the annual baseline.

Table A3: **Parameters Estimated with SMM - annual vs quarterly**

Parameter	Description	Annual	Quarterly
σ	Elasticity of substitution	6.831	7.954
α	Elasticity of output with respect to labor	0.195	0.218
β	Bargaining power of workers	0.457	0.463
θ	Elasticity of the matching function	1.875	0.839
δ	Exogenous exit hazard	0.046	0.018
c_h	Scalar, vacancy cost function	0.696	0.466
λ_1	Convexity, vacancy cost function	2.085	2.384
λ_2	Scale effect, vacancy cost function	0.302	0.307
b	Value of home production	0.403	0.280
l_e	Initial size of entering firms	6.581	2.560
c_p	Fixed cost of operating	10.006	9.882
c_x	Fixed exporting cost	100.23	65.674
c_e	Entry cost for new firms	25.646	23.665
σ_z	Standard deviation of the z process	0.135	0.0797

The other major change in parameter values is the decrease in matching function elasticity θ from 1.875 to 0.839. This compensates for the fact that, other things equal, switching to a quarterly frequency would have increased labor market tightness as workers enjoyed more opportunities to match with firms. In turn, this would have made it more difficult for firms hire, and thus shifted the simulated firm size distribution leftward. Dropping θ improves the ability of firms to meet workers over the relevant range of (U, V) values, and thus prevents this from occurring. Other parameter values such as exogenous exit rate δ and the initial firm size l_e drop in proportion to the change in model frequency.

Table A5 addresses the main question of interest: how robust are the policy experiments in the paper to the unit of time used in the model? That is, it redoes Table 4 using the quarterly version of our model. Note that here, as in the paper, the "reforms and globalization exercise" (3rd column) is based on a level of iceberg costs τ_c that induces the observed post-reform fraction of firms that export. However, for the quarterly version, a smaller reduction in τ_c is needed to hit this target.

Comparing these simulation outcomes with the annual baseline in the manuscript reveals several patterns. First, qualitative responses to tariffs, iceberg costs and firing costs (τ_a, τ_c, and c_f) are robust to the model's periodicity. Job turnover, firm-level wage dispersion and industrial sector unemployment all increase in response to reductions in the "reforms and globalization" experiment (3rd column). Also, as in the annual version of the model, the size distribution shifts rightward while there is a sizeable drop in the mass of firms. However, the quantitative responses of job turnover and wage dispersion are somewhat different. In our annual model, the "reforms and globalization" experiment increased job turnover by about 12 percent, leaving worker-level log wage dispersion relatively stable. In the quarterly model, the same experiment increased job turnover by only 4 percent, but increased the standard deviation of log wages by 9 percent. These contrasts reflect the shifts in parameter estimates described above. With a smaller home production payoff, b, wages are more sensitive (percentagewise) to firm characteristics (z, l). On the other hand, a lower matching function elasticity makes job finding and fill rates less responsive to changes in aggregate labor market conditions, muting the increase in job turnover.

Table A4: **Data-based versus Simulated Statistics - annual vs quarterly**

Moment	**Data**	**Annual**	**Quarterly**	**Size Distribution**	**Data**	**Annual**	**Quarterly**
$E(\ln G_t)$	5.442	5.253	5.199	20th percentile cutoff	14.617	15.585	16.787
$E(\ln l_t)$	3.622	3.636	3.587	40th percentile cutoff	24.010	25.773	26.024
$E(\mathcal{I}_t^x)$	0.117	0.108	0.126	60th percentile cutoff	41.502	41.432	44.707
$var(\ln G_t)$	2.807	3.329	5.348	80th percentile cutoff	90.108	79.109	75.48
$cov(\ln G_t, \ln l_t)$	1.573	1.788	2.681	**Firm Growth Rates**			
$var(\ln l_t)$	1.271	1.219	1.679	<20th percentile	1.421	1.234	1.477
$cov(\ln G_t, \mathcal{I}_t^x)$	0.230	0.251	0.320	20th-40th percentile	0.255	0.271	0.427
$cov(\ln l_t, \mathcal{I}_t^x)$	0.152	0.160	0.194	40th-60th percentile	0.209	0.183	0.259
$cov(\ln G_t, \ln G_{t+1})$	2.702	2.196	-0.254	60th-80th percentile	0.184	0.151	0.167
$cov(\ln G_t, \ln l_{t+1})$	1.538	1.556	3.524	**Aggregate Turnover/**			
$cov(\ln G_t, \mathcal{I}_{t+1}^x)$	0.225	0.278	0.621	**Wage Dispersion**			
$cov(\ln l_t, \ln G_{t+1})$	1.543	1.394	1.653	Firm exit rate	0.108	0.120	0.092
$cov(\ln l_t, \ln l_{t+1})$	1.214	1.161	1.276	Job turnover	0.198	0.240	0.315
$cov(\ln l_t, \mathcal{I}_{t+1}^x)$	0.152	0.185	0.209	Std. dev. of log wages	0.461	0.426	0.471
$cov(\mathcal{I}_t^x, \ln G_{t+1})$	0.220	0.279	0.285	**Olley-Pakes Statistics**			
$cov(\mathcal{I}_t^x, \ln l_{t+1})$	0.149	0.201	0.203	$(1-\alpha)\left(\frac{\sigma-1}{\sigma}\right)$	0.685	0.687	0.684
$cov(\mathcal{I}_t^x, \mathcal{I}_{t+1}^x)$	0.089	0.073	0.089	d_F	0.090	0.094	0.094

Table A5: **Effects of Reforms and Globalization under Quarterly Model Period**

	Baseline	Reforms	Reforms & Globalization
τ_a (ad valorem tariff rate)	1.21	1.11	1.11
c_f (firing cost)	0.6	0.3	0.3
τ_c (iceberg trade cost)	2.5	2.5	2.2
Size Distribution			
20th percentile	17	17	19
40th percentile	26	27	32
60th percentile	45	48	61
80th percentile	75	81	116
Average firm size	31	34	46
Firm Growth Rates			
<20th percentile	1.48	1.47	1.51
20th-40th percentile	0.43	0.43	0.52
40th-60th percentile	0.26	0.27	0.34
60th-80th percentile	0.17	0.16	0.21
Aggregates			
% of firms exporting	1	1.44	2.73
Revenue share of exports	1	1.45	2.73
Exit rate	1	0.98	1.04
Job turnover	1	0.99	1.04
Mass of firms	1	0.91	0.67
Unemployment rate in the industrial sector	1	1	1.13
Industrial share of employment	1	0.98	0.98
Standard deviation of log wages (firms)	1	1.05	1.12
Standard deviation of log wages (workers)	1	1.06	1.09
Log 90-10 wage ratio (firms)	1	1.05	1.14
Log 90-10 wage ratio (workers)	1	1.07	1.12
Standard deviation of workers' value (J)	1	1.14	1.09
Log 90-10 ratio of workers' value (J)	1	1.16	1.12
Real income	1	1.02	1.21

Note: Aggregate statistics in the bottom panel are normalized by their baseline levels.