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Trade liberalization and labor demand elasticities: evidence from Turkey

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Abstract

In the recent debate over the impact of trade reform on factor markets, it has been argued that trade liberalization will lead to an increase in labor-demand elasticities — thus placing labor markets under increased pressure. Using Turkish plant-level data spanning the course of a dramatic trade liberalization, we test this idea. However, we are unable to find any empirical support for this supposed theoretical link: in most of the industries we consider, we cannot reject the hypothesis of no relationship between trade openness and labor-demand elasticities.

Keywords: Trade reforms; Imperfect competition; Factor demand elasticities; Turkey

JEL classification: F02; F12; F13; F15

1. Introduction

The finding of increased wage inequality since the 1970s in the US and other countries has generated intense controversy amongst both economists and policy makers regarding its causes. While no consensus exists on the driving forces behind these changes, their coincidence with increased trade openness over this period has led to a revival of interest in the supposed links between trade and labor

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markets. These linkages are now being re-examined and analyzed in a large and growing theoretical and empirical literature.[1] While the primary focus of this research has appropriately been on the direct impact of greater openness on employment, wages and particularly on wage inequality (between high and low-skilled workers),[2] the literature (for instance, Rodrik (1997) and Slaughter (1997)) has emphasized yet another linkage between openness and labor demand: the possibility, particularly in imperfectly competitive contexts, for the elasticity of demand for labor to be higher with greater openness. As Slaughter (1997) has pointed out, the link between factor demand elasticities and product market elasticities is directly established through Hicks' well known 'fundamental law of factor demand' which implies that ''the demand for anything is likely to be more elastic, the more elastic is the demand for any further thing which it contributes to produce''. Since product market elasticities are likely to rise with trade liberalization, this implies that, with greater trade openness, we should see an increase in labor demand elasticities as well.[3]

Why do rising labor demand elasticities matter? As Rodrik (1997) notes, rising elasticities have three important consequences. First, they shift the wage or employment incidence of non-wage labor costs towards labor and away from employers. Second, higher elasticities trigger more volatile responses of wages and employment to any exogenous shocks to labor demand. Third, higher elasticities shift bargaining power over rent distribution in firms enjoying extra-normal profits away from labor and towards capital. Thus, the overall point is that moves towards openness could put labor markets under greater 'pressure' — an outcome that many would consider to be socially undesirable.

This paper attempts to investigate the link between between trade openness and factor demand elasticities empirically.[4] We start by specifying an econometrically implementable theoretical model of a firm operating in an imperfectly competitive context and derive predictions about the implications of changes in trade policy for labor demand elasticities. This is then tested using data from the Turkish

[1] In the recent literature, early contributions include Revenga (1992), Bhagwati and Dehejia (1994), Lawrence and Slaughter (1993), Sachs and Schatz (1994), Freeman and Katz (1994) and Currie and Harrison (1997).

[2] In particular, the implications of the Stolper–Samuelson theoerm (which ultimately predicts in the typical $2\times2\times2$ context that with greater openness a country see an improvement in the rewards to its abundant factors) have received much empirical attention.

[3] See Hammermesh (1993) and Slaughter (1997) for a more detailed discussion. It should be emphasized also that the argument just stated can be made directly only in a partial equilibrium context. For a critical examination of this linkage between openness and labor demand elasticities in a general equilibrium context, see Panagariya (2000).

[4] There have, of course, been other studies of linkages between trade policy and labor markets: Revenga (1992), Harrison and Revenga (1995), Kambhampati et al. (1997) and Currie and Harrison (1997) are recent examples. By and large, each of these papers focuses on the relationship between trade policy and employment and wage levels rather than on elasticities. Slaughter (1997), on the other hand, uses industry-level US data to investigate the link between variations in openness in the recent decades and labor demand elasticities.

manufacturing sector from a period when there were large scale changes in the level of trade protection (specifically, the trade reforms of 1984).[5] This data set provides a unique opportunity to test the impact of trade policy on labor demand elasticities. It has several appealing features: First, the data span years in which the trade policy change was really quite dramatic. Thus, if it is the case that greater openness generally results in larger factor demand elasticities, we should expect to see its effects after this particular reform.[6] Second, the data are of higher frequency than is typical: since manufacturing censuses are typically only conducted once every 3 or 5 years in most countries, one is usually only able to obtain two or three observations per firm in about a decade — a period over which there may typically be multiple shifts in trade policy. Thus, in most cases, extracting information about the impact of a single shift in trade policy on firms is difficult. Our data, however, are annual, thus mitigating this concern. Third, this data set has already been used previously by Levinsohn (1993) to examine the impact of the trade reforms on the competitive behavior of firms. It, thus, permits an industry by industry comparison and conciliation of product market impact with impact on labor markets.

Our analysis suggests that the linkage between greater trade openness and labor demand elasticities as suggested by the theory may be empirically quite weak: in the vast majority of the industries we considered, we are unable to reject the hypothesis of no relationship between these variables. As we discuss in detail in the paper, this finding remains robust to changes in the type of labor considered (all production workers, overtime labor, externally contracted labor, female labor, etc.) and quite robust to changes in specification as well.

The rest of the paper is structured as follows. Section 2 describes the Turkish reforms briefly. Section 3 outlines the basic theoretical structure and the derives the estimating equation. Section 4 describes the data. Section 5 discusses the primary econometric concerns and presents the econometric results. Section 6 discusses how we can reconcile our results with the findings of Levinsohn (1993) who finds significant impact of the trade reforms in product markets. Section 7 concludes.

2. The Turkish reforms

As numerous studies have noted,[7] prior to the 1980s, the import competing sector in Turkey received an extraordinarily high level of protection. The Turkish

[5] This same episode, using the same exact data set, was examined in the classic paper by Levinsohn (1993) which looked at the impact of trade reforms on industry markups and found strong support in the data for the hypothesis that greater openness leads domestic firms to behave more competitively.

[6] In this regard, our study is set in a better context than the one examined by Slaughter (1997) whose focus instead is on US trade policy changes over the last decades which were rather less pronounced and more difficult to measure.

[7] See, in particular, Krueger (1974) and Yagci (1984).

trade regime, comprising of a complicated combination of quotas, tariffs, import licensing procedures and foreign exchange regulations was considered virtually peerless in its complexity and in the degree of protection it offered to its import competing sector.[8] Indeed, as Levinsohn (1993) notes, ''trade policy was a protectionist's dream'': the average tariff in 1981 was estimated to be 49% with the general pattern of tariffs being one where rates were lowest on intermediate inputs and highest on finished products. Further, for over half the products, tariff equivalents of non-tariff barriers were estimated to be over 100% (Krueger, 1974).

Significant import liberalization measures were announced in December 1983 and implemented soon after.[9] The 1984 import liberalization program significantly reduced both tariff and non-tariff barriers. The previous regime was replaced by one in which quotas were expanded, fewer goods were placed on the 'prohibitive list' and the proportion of imports covered dropped significantly. Almost all consumer goods were liberalized. Tariffs were reduced significantly as well (Krueger and Turan, 1993).

Table 1 (reproduced from Levinsohn (1993), itself partially adapted from Baysan and Blitzer (1990)), notes the three-digit ISIC industries in which, in 1984, Turkey was a net importer. Table 1 also lists import levels in these industries in 1984 and the estimated changes in protection following the reforms. In all but two industries, protection was estimated to have fallen in 1984. It is the impact of these reductions in trade protection on labor markets that this paper studies.

Table 1
Importing Turkish industries: categories, import volumes, protection and markup changes

ISIC code	Category title	Protection change	1984 Imports US $1000s	Markup change
341	Paper	Decrease	101 109	Increase
351	Industrial chemicals	Decrease	1 506 921	No change
352	Other chemical products	Decrease	150 777	Decrease
361	Pottery	Decrease	15 633	Decrease
372	Non-ferrous metals	Decrease	218 484	Decrease
381	Metal products	Decrease	173 130	Increase
382	Non-electrical machinery	Decrease	1 385 094	Decrease
383	Electrical machinery	Decrease	590 233	Decrease
384	Transport equipment	Increase	606 134	Increase
385	Scientific equipment	Increase	130 431	Increase

[8] It had, nevertheless, quite good company — the Indian trade regime, for instance.

[9] The initial move towards liberalization of the economy in general, had started by 1980 itself. Fortunately for this study (since the first year of available data is 1983), significant liberalization of imports only started later.

3. Theory and the estimation framework

To demonstrate theoretically how changes in trade policy resulting in greater product market competition and larger product market elasticities could work their way to larger factor demand elasticities, and to establish theoretical underpinnings for the empirical work to follow, we work with a model of monopolistic competition, where each firm faces its own less than infinitely elastic demand curve and where there is assumed to be no strategic interaction between firms.[10] Thus, any firm i in industry j is assumed to face an inverse demand curve of the type:

$$P_{ij} = \theta \bar{P}_j Q_{ij}^{-1/\in} \tag{1}$$

where P_{ij} denotes own price, $\bar{P}_j$ denotes industry average price, θ is a scaling factor, Q_{ij} denotes firm output and $\in$ denotes the (constant) price elasticity of demand. The production function is assumed to be of a Cobb–Douglas type (in variable inputs), and is given by:

$$Q_{ij} = \prod_{k=1}^{n} V_{kij}^{\alpha_k} \tag{2}$$

where V_{kij} denotes the kth input in use here.

The firm is assumed to face infinitely elastic factor supplies, i.e., it takes factor prices as given. Partially differentiating profits with respect to the lth input and equating it to zero gives us the following first order condition:

$$\theta \bar{P}_j Q_{ij}^{1-1/\in} \left(1 - \frac{1}{\in}\right) \alpha_l V_{lij}^{-1} = w_l \tag{3}$$

where w_l denotes the price of the lth input. In log form, (3) can be rewritten as:

$$\ln V_{lij} = \gamma_0 + \gamma_1 \ln\left(\frac{w_l}{\bar{P}_j}\right) + \sum_{k \neq l} \gamma_k \ln V_{kij} \tag{4}$$

where the γ values are each a function of $\in$. Substituting the FOC for other inputs back into (4) above, we get:

$$\ln V_{lij} = \delta_0 + \sum_{k=1}^{n} \delta_k \ln(w_k/\bar{P}_j) \tag{5}$$

which is a demand function for the lth input expressed in terms of the variables the firm takes as exogenous — factor prices and the industry average price. The

[10] This approximates a situation in which there are a large number of varieties and each firm is an infinitesimal player but has some power over the pricing of its product.

coefficients in the above equation are functions of $\in$. The own price elasticity of an input is given by:

$$\frac{\partial \ln V_{lij}}{\partial \ln(w_l/\bar{P}_j)} = \delta_l = \frac{-\left[1-\left(1-\frac{1}{\in}\right)\left(\sum_{k\neq l}\alpha_k\right)\right]}{\left[1-\left(1-\frac{1}{\in}\right)\left(\sum_{k=1}^{n}\alpha_k\right)\right]} \tag{6}$$

The derivative of the absolute value of the own price elasticity of input demand with respect to the product demand elasticity is given by

$$\frac{\partial|\delta_l|}{\partial \in} = \frac{\alpha_l}{\in^2\left[1-\left(1-\frac{1}{\in}\right)\left(\sum_{k=1}^{n}\alpha_k\right)\right]^2} > 0 \tag{7}$$

which shows that the magnitude of the own price elasticity of factor demand is increasing in the product demand elasticity (thus proving that the effects predicted by Hick's Law are relevant in this context). The cross-price elasticity of demand for inputs may be analogously derived and is given by:

$$\frac{\partial \ln V_{lij}}{\partial \ln(w_s/\bar{P}_j)} = \delta_s, \quad s \neq l \tag{8}$$

We now specialize to the case where the only inputs are labor, capital, materials and fuel. Additionally, we let w, r, m and f denote the logs of the wage rate, the rental rate, materials price and the fuel price, respectively, each deflated by the industry-level average output price. Finally, we focus our attention on the demand for labor alone (as opposed to that for other factor inputs).[11] The labor demand function is then given by

$$l_{ijt} = \delta_0 + \delta_w w_{ijt} + \delta_r r_{ijt} + \delta_m m_{ijt} + \delta_f f_{ijt} \tag{9}$$

where l is the log of labor demanded. Thus, our final estimating equation is

$$l_{ijt} = \delta_0 + \delta_w w_{ijt} + \delta_r r_{ijt} + \delta_m m_{ijt} + \delta_f f_{ijt} + e_{ijt} \tag{10}$$

where the error term e_{ijt} allows for random shocks to affect the firm's demand for labor.

[11]The theory clearly applies to all input factors. However, we focus only on labor demand since any effort at estimating demand equations for other factors would be frustrated by the fact that firm level data on the use of the other factors and/or factor prices is not available.

4. Data

As we have already noted, the data used in this paper are from the Turkish manufacturing census for 10 different three-digit ISIC industries and span the years 1983–1986. The data are annual and cover all plants in the greater Istanbul area. Since the manufacturing sector in Turkey is heavily concentrated around the Istanbul area these provide a fairly comprehensive coverage of Turkish Industry. Table 1 describes the import competing industries that we consider in this paper. The industries range from paper to industrial chemicals to electrical machinery and manufacturing of scientific equipment. Table 1 also lists the volume of import by industry, estimates of (directions of) protection change (which is the sum of changes in protection in tariff and non tariff protection) and Levinsohn's (1993) estimates of (directions of) markup changes in these industries.

As Table 1 indicates, eight of the 10 industries saw a reduction in protection with the trade reforms and the rest two saw an increase. Levinsohn's estimates of markup change are strongly correlated with these changes in the level of protection: in seven out of 10 cases, the estimated change in markups is directly related to changes in the level of protection. Two out of these 10 cases, go in the other direction, however, and for industrial chemicals, Levinsohn reports no change in markup.

Our interest in this paper is in labor market impact. To estimate (10), we need data on input prices faced by the firm and labor employed. Our basic measure of labor is the same as the one used by Levinsohn (1993). We focus on production labor to obtain which we sum male and female high level and medium level technical personnel, male and female foremen and male and female workers. For robustness, we also estimate (10) using other categories of labor, specifically, female workers, contract workers, and overtime labor.

The nominal firm level wage is obtained by dividing the payroll corresponding to production workers by the above sum. The data set provides information directly on three-digit industry-specific raw materials input price indices and on fuel prices. Economy-wide annual interest rate data (borrowing costs) were used as rental costs.[12]

5. Econometric issues and estimation results

Eq. (10) derived in the previous section is our basic estimating equation. We estimate (10) separately for each industry (although results using data pooled across industries are presented and discussed later in this section as well). To take into account within-industry firm heterogeneity, we model firm-specific effects that

[12] We find, however, as does Levinsohn (1993) that the results are mostly invariant to the assumptions regarding interest rates. This is discussed in more detail in the next section.

are constant over time by including firm-specific intercepts in (10). Both 'fixed effects' and the 'random effects' specifications are estimated. To capture the effect of change in trade policy on the parameters in (10), we include intercept and interactive trade reform dummies which take the value of one for the post liberalization period. We also experiment with year-specific intercept dummies (in place of the reform intercept dummy) to capture year-specific shocks common to all firms in an industry.

5.1. Industry-specific own price labor demand elasticities and post-reform changes

Our estimates of labor demand elasticities and their changes in each of the 10 industries, under the fixed effects and random effects specifications, are presented in Table 2. The vast majority of the estimated elasticities (δ_w) lie within the range of -0.15 to -0.75. Thus, these fall well within what Hammermesh (1993) has identified as being a reasonable range of values for labor demand elasticities. In eight out of 10 cases, under both fixed effects and random effects specifications, the elasticity estimates are quite tightly estimated.

The parameter of particular interest here is elasticity change, i.e., the parameter corresponding to the wage variable interacted with the liberalization dummy — $\Delta\delta_w$ in Table 2. As this table indicates, our estimates of the changes in labor demand elasticities are small in magnitude and largely insignificant. In seven out of 10 cases, under both the fixed effects and random effects specifications, the null hypothesis that the change in elasticity after the reforms is zero cannot be rejected at the 5% level or indeed in most cases at even a higher level of significance. The three industries where the null hypothesis of no elasticity change is rejected are Metal Products (381), Non-Electrical Machinery (382) and Electrical Machinery

Table 2
Own price labor demand elasticity estimates: fixed effects and random effects[a]

ISIC code	Fixed effects		Random effects	
	δ_w	$\Delta\delta_w$	δ_w	$\Delta\delta_w$
341	−0.66 (0.10)	0.04 (0.11)	−0.51 (0.10)	0.07 (0.12)
351	0.02 (0.17)	−0.02 (0.18)	0.014 (0.02)	−0.02 (0.18)
352	−0.43 (0.05)	0.02 (0.05)	−0.34 (0.06)	0.01 (0.06)
361	−1.03 (0.11)	0.18 (0.15)	−1.00 (0.12)	0.20 (0.15)
372	−0.94 (0.05)	0.11 (0.10)	−0.9 (0.06)	−0.15 (0.10)
381	−0.63 (0.05)	−0.10 (0.05)	−0.53 (0.05)	−0.11 (0.05)
382	−0.57 (0.06)	−0.12 (0.07)	−0.49 (0.06)	−0.14 (0.06)
383	−0.56 (0.06)	−0.14 (0.06)	−0.42 (0.06)	−0.18 (0.06)
384	−0.53 (0.10)	−0.1 (0.09)	−0.24 (0.11)	−0.14 (0.11)
385	−0.33 (0.09)	−0.06 (0.10)	−0.06 (0.10)	−0.04 (0.10)

[a] Note: figures in parentheses are standard errors.

(383). In these cases the $\Delta\delta_w$ estimate is negative, implying that the absolute value of the own price labor demand elasticity goes up. However, in one out of these three industries, namely, Metal Products (381), Levinsohn actually found an increase in markup implying a reduction in the product demand elasticity perceived by plants in this industry. Overall then, it appears that in our data industries' labor demand elasticities are subject to friction and do not respond to changes in openness as predicted by the theory.

Alternative specifications were attempted as well: in order to take into account the possible fixity of capital in the short run, (10) was estimated by dropping the terms corresponding to rental rate of capital. Time-specific intercept dummies (in place of the reform dummy) were included (in addition to the firm-specific effects) in our regressions. The results remain more or less the same with the estimates of own price labor demand elasticities, their changes and the associated standard errors all changing only negligibly.[13]

Several issues regarding the validity of the estimation framework and the interpretation of the results arise. First is the familiar issue of possible simultaneity and correlation between the error term and the right-hand side variables. Since both labor demand and labor supply depend upon the wage, shocks to the labor demand will result in shocks to the wage.[14] Thus the wage and the disturbance term in our estimating equation may be correlated, thereby raising the possibility of a bias in our estimates.[15] Another source of correlation between the wage variable and the disturbance term is measurement error. Given that the firm level wage rate is determined by dividing the wage bill by the number of workers, the measurement error in the labor variable also contaminates the wage variable.

As regards the endogeneity of the wage to changes in labor demand, the identifying assumption here clearly is that labor supplies facing each firm are perfectly elastic, i.e., that shifts in the labor supply curve (resulting in changes in wages) trace out the labor demand schedule and shocks to the labor demand do not affect wages. As Hammermesh (1993) notes, the suitability of this identifying assumption rests on the degree of disaggregation of the data. Since the data we use are plant level, thus quite heavily disaggregated, we do not consider this to be a serious issue here. In addition, all the plants we are looking at are in a common geographical location — the Greater Istanbul area. There are about 600 plants in our sample. Given that so many plants and almost all of Turkish manufacturing

[13]These results are available from the authors upon request.

[14]To the extent that, say, aggregate demand or productivity shocks increase product demand and raise labor demand and increase wages (or any other factor prices for that matter) at the same time, the elasticity estimates delivered from (6) would be biased due to a correlation between the error term and the right had side variable.

[15]In a well known contribution, Nickell and Symons (1990) have argued that the identification problem does not really exist anyway since labor supply and labor demand really depend upon two quite different real wages — one deflated by the producer price and one by the consumer price index. Thus using the appropriate real wage implies that simultaneity should not be a real problem.

industries are located in and around one city (Istanbul), it is highly improbable that any one buyer of factor inputs will have any market power in these factor markets. Thus, the plausibility of our identifying assumption is certainly greater than in most studies of this nature that use industry level data instead. Furthermore, as stated above, we find that introducing time-specific dummies in addition to firm-specific effects does not change our elasticity and elasticity change estimates. The results with both these kinds of effects are negligibly different from those presented in Table 2. This is reassuring since any aggregate demand or productivity shocks (which may simultaneously move labor demand and wage as noted earlier) are thus accounted for — taking care of the bulk of this endogeneity problem.

While instrumental variables estimation is perhaps a more satisfactory approach, it proves a little less feasible in this context: other than lagged endogenous variables, we have no variables in the data set that we may regard as being exogenous. As Levinsohn (1993) has already noted in his use of this data, using lagged variables as instruments is problematic due to the short length of our panel (4 years) — the number of observations prior to the reforms is cut by half and overall the number of observations is down by a fourth even when we only use single year lags. Estimating (10) using lagged variables as IVs resulted in highly insignificant (and sometimes meaningless) estimates of the parameters of interest. One option that presents itself then is the pooling of data across industries to use lagged variables as instruments. These results (which turn out to be qualitatively the same as our uninstrumented fixed and random effects results) are presented in Table 5 and are discussed later along with other results from the pooled sample.

Finally, we must note, as does Slaughter (1997), that even though a correlation between the wage and the error term will bias the elasticity estimates, there is no reason to expect the post-liberalization elasticity change estimates to be biased one way or the other. More precisely, there is no reason to expect the bias in the labor demand elasticity estimate to be different in one regime (post- or pre-reform) than in the other. To see whether this may have been driven simply due to the problem that we just noted, we conducted Monte Carlo Simulations which, given the data, look for a biased estimate of the change in elasticity due to a (programmed) correlation between the right-hand side variable and the error term. For each assumed correlation we examined, we find this not to be the case (See Table 7 for the Monte Carlo results and Appendix A for a description of procedure). That is, the Monte Carlo simulations do not suggest a bias that attenuates the estimate of the change in elasticity to zero as we find. Besides, our IV estimates with plants from all industries pooled together confirm our beliefs regarding the lack of bias in the elasticity change estimate $\Delta\delta_w$.

A second issue concerns that of timing and lagged responses. It is assumed in our estimation of (10) that firm demand responses to changes in wage rates occurred without lags. As Hammermesh (1993) has noted, much of the adjustment in firm labor demand takes place within 6 months to 1 year. Thus, given that our data are annual, this is not a serious problem.

A third issue is that of constancy of parameters across firms within an industry — or alternately of our implicit estimation assumption that firms within an industry have identical wage elasticities. The data indicate that the measured (average) wage is quite different across firms within the same industry. We believe that this reflects unobserved differences in (average) worker quality across firms or to a smaller extent due to differences in the number of hours on the job put in by workers in different firms. Given such differences, it would be reasonable to expect that labor demand itself could be somewhat different across firms within an industry. Of course, to the extent that the differences are simply in levels and are fixed over time — our firm specific intercept should take care of the problem. However, one may expect the slopes to be different across firms within an industry as well. It is, nevertheless, infeasible to estimate firm specific elasticities and their changes for each firm when we only have four observations per firm. In order to address this issue, we experimented with a random coefficients (Hildreth–Houck) specification where the parameter estimates are firm-specific but assumed to be drawn from a distribution that is common across firms within a given industry. The results remain the same qualitatively, i.e., the estimated changes in elasticities after the trade liberalization continue to be insignificantly different form zero in all cases.[16]

5.2. Cross price elasticities of labor demand

Cross price elasticities (of labor demand) and their changes following the trade reform were not estimated with great precision. Hardly any of the estimates were significant at the 5% level or even at higher levels. Fixed effects and random effects estimates of the cross materials price labor demand elasticity estimates are presented in Table 3. We are not greatly surprised by these estimates and their lack of significance — with few exceptions, the literature has traditionally found cross price elasticities to be rather difficult to estimate. Few examples of success can be found. This is additionally difficult in our case since in, our data, factor prices for materials and fuel are only specific to the industry (i.e., are not available at the level of the firm) and thus they exhibit insufficient variation over the sample.

5.3. Labor demand elasticities (and changes) by worker type

For robustness (and also independent interest in variations in labor demand elasticities across worker types), we estimate (10) by considering the demand for female workers, contract workers and overtime workers separately. The elasticities again are quite tightly estimated (see Table 4). However, as expected, their values are higher in magnitude than the ones for overall labor as substitution possibilities are higher when we look at specific kinds of labor than in the case of labor in

[16] The results with random coefficients are not presented in this paper. However, they can be obtained from the authors on request.

Table 3
Cross materials price labor demand elasticity estimates: fixed effects and random effects[a]

ISIC code	Fixed effects		Random effects	
	δ_m	$\Delta\delta_m$	δ_m	$\Delta\delta_m$
341	−2.76 (2.04)	7.43 (2.73)	−2.17 (2.22)	7.02 (2.96)
351	0.75 (2.98)	0.17 (3.27)	1.12 (2.95)	−0.18 (3.23)
352	−11.14 (3.31)	9.07 (3.67)	−10.71 (3.59)	8.58 (3.97)
361	−2.60 (1.52)	–	−2.48 (1.57)	–
372	2.37 (1.38)	0.19 (1.94)	2.48 (1.50)	0.10 (2.12)
381	−11.88 (11.71)	9.31 (11.74)	−8.46 (12.30)	5.87 (12.33)
382	1.30 (1.13)	−4.57 (1.31)	1.18 (1.19)	−4.30 (1.37)
383	1.55 (1.54)	–	1.62 (1.69)	–
384	29.15 (33.61)	28.58 (33.63)	27.70 (37.93)	−27.00 (37.96)
385	−3.48 (2.88)	5.29 (3.69)	−3.72 (2.98)	5.44 (3.81)

[a] Note: figures in parentheses are standard errors.

general. For female workers, most elasticity estimates are between −1 and −2, while for contract labor the estimates can go as high as −3.65. For overtime workers, labor demand seems to be inelastic. Changes in elasticity are again mostly insignificant following the reforms: significant negative values for $\Delta\delta_w$ are produced in very few cases — Electrical Machinery (383) and Scientific Equipment (385) for female workers and Transport Equipment (384) for overtime workers. It should be noted in this context that both Transport Equipment and Scientific Equipment experienced an increase in protection in the post-reform period and Levinsohn's estimates actually show an increase in markups, implying a fall in the firm's perceived product demand elasticity. Thus, considering disaggregated categories of labor does not alter the inference that we arrived at

Table 4
Own price labor demand elasticities by worker type: fixed effects estimates[a]

ISIC code	Female workers		Contract workers		Overtime workers	
	δ_w	$\Delta\delta_w$	δ_w	$\Delta\delta_w$	δ_w	$\Delta\delta_w$
341	−1.56 (0.28)	−0.30 (0.26)	−0.76 (1.13)	−0.09 (1.09)	−1.09 (0.42)	0.25 (0.59)
351	−0.64 (0.42)	0.52 (0.52)	−3.65 (1.44)	1.92 (1.23)	−0.51 (0.32)	−0.52 (0.36)
352	−1.42 (0.09)	0.17 (0.08)	−0.85 (0.36)	0.29 (0.36)	−0.35 (0.22)	−0.15 (0.22)
361	−1.89 (0.64)	0.35 (0.41)	−2.50 (3.33)	0.97 (2.09)	−0.66 (0.27)	−0.51 (0.29)
372	−1.23 (0.55)	0.46 (0.27)	−1.17 (0.56)	0.23 (0.63)	−1.32 (0.16)	−0.06 (0.44)
381	−1.33 (0.18)	−0.18 (0.16)	−1.19 (0.24)	0.32 (0.23)	−0.85 (0.26)	−0.23 (0.26)
382	−2.00 (0.34)	1.05 (0.37)	−0.99 (0.20)	−0.30 (0.21)	−0.70 (0.17)	0.40 (0.18)
383	−0.96 (0.15)	−0.33 (0.15)	−1.09 (0.28)	−0.14 (0.27)	−0.51 (0.19)	0.10 (0.20)
384	−0.16 (0.30)	−0.14 (0.28)	−2.26 (0.31)	−0.01 (0.26)	−0.69 (0.24)	−0.52 (0.22)
385	−0.50 (0.17)	−0.44 (0.20)	−0.68 (0.84)	0.24 (0.99)	0.24 (1.76)	0.25 (0.73)

[a] Note: Figures in parentheses are standard errors.) Random effects estimates are qualitatively similar and can be obtained from the authors on request.

earlier (from the aggregated data) that labor demand elasticities seem to be unresponsive to openness.

5.4. Results of estimation using data pooled across industries

Our penultimate set of regressions used data pooled across all the 10 industries. The results are presented in Table 5. As noted earlier, this was done with the intention of using instrumental variables techniques to control for any remaining simultaneity problems in the framework. Pooling of the data also allows us to use variations in cross industry changes in the protection levels and in import penetration ratios as proxies for the reform dummies.[17]

We start first with estimates of (10) using pooled data and using lagged right-hand side variables as instruments (Regression A). The IV estimates for the pooled sample gives us an elasticity estimate that is not as precisely estimated (although the point estimate appears to be in a reasonable quantitative range) and again we cannot reject the null that the trade reform did not have any effect on the

Table 5
Regressions using data pooled over all industries

	Regression A			
	Instrumental variable estimates with reform dummy interactions			
	δ_w	-0.53 (0.57)		
	$\Delta\delta_w$	-0.02 (0.03)		
	Regression B		Regression C	
	With actual tariff Interactions		With import penetration interactions	
Fixed effects	δ_w	-0.55 (0.02)	δ_w	-0.57 (-0.01)
	$\partial\delta_w/\partial t$	-0.17 (0.07)	$\partial\delta_w/\partial(\textit{imp-pen})$	-0.02 (0.03)

[17] We are most grateful to a referee for this suggestion but should also note some caveats associated with pooling data across industries in this context: First, the implicit assumption associated with pooling is that the economic response to greater openness would be similar in magnitude across industries if the degree to which openness increased was similar across industries. However, there is no reason to expect this. To see this, note that in our derivation of (10), the labor demand elasticity is a function of both product demand elasticities and the Cobb–Douglas production coefficients. Neither need be similar in magnitude across industries. Separately from this, given the fact that our estimates of labor demand elasticities are not similar across industries (as Table 2 indicates) and the fact that the estimated standard errors of the error terms of industry-specific regressions vary a great deal across industries, imposing common coefficients in a pooled regression raises the problem of bias in our estimates (as Greene, 1993, Chapter 7, has noted).

labor demand elasticity. The point estimate of the change in elasticity is small in magnitude indicating that it is economically insignificant as well.

Regression B uses actual tariff rate (in place of the reform dummy) interactions instead. The coefficient of the cross product of t and the log of wage (deflated by industry price) gives us $\partial\delta_w/\partial t$, the derivative of the labor demand elasticity with respect to the tariff rate. The labor demand elasticity at zero tariffs, δ_w, is estimated to be -0.55 and is significant at the 1% level. The estimate for $\partial\delta_w/\partial t$ is -0.17 and is significant at the 5% level. Surprisingly and counter-intuitively, the sign of this coefficient shows that higher tariffs are associated with higher magnitudes of labor demand elasticities.

Regression C uses import-penetration interactions. The coefficient of the cross product of *imp-pen* and the log of wage (deflated by industry price) gives us $\partial\delta_w/\partial(\textit{imp-pen})$, the derivative of the labor demand elasticity with respect to the import-penetration ratio. The labor demand elasticity at zero import penetration ratio, δ_w, is significantly estimated to be -0.57. The estimate for $\partial\delta_w/\partial(\textit{imp-pen})$ is -0.02 and is insignificant.

6. Reconciliation with the Levinsohn (1993) results

Our finding that greater trade openness did not lead to greater labor demand elasticities in Turkey at first thought seems somewhat inconsistent with Levinsohn's (1993) finding that greater openness did lead to reduced markups (just as theory would predict). This is all the more puzzling since both the markup equation estimated by Levinsohn and the elasticity equation that we estimate instead follow from the same set of first-order conditions for profit maximization for a firm operating in an imperfectly competitive context.

To see this more clearly, note that in exact correspondence to the first-order conditions (relating marginal revenue to marginal cost) of the firm's maximization problem in our case (i.e., Eq. (3)), Levinsohn's paper derives the following first-order conditions:

$$\frac{\dfrac{\partial q_{ijt}}{\partial[labor]_{ijt}}}{\left[\dfrac{w_{ijt}}{p_{jt}}\right]} = \frac{\dfrac{\partial q_{ijt}}{\partial[capital]_{ijt}}}{\left[\dfrac{r_{ijt}}{p_{jt}}\right]} = \frac{\dfrac{\partial q_{ijt}}{\partial[materials]_{ijt}}}{\left[\dfrac{m_{ijt}}{p_{jt}}\right]} = \frac{\dfrac{\partial q_{ijt}}{\partial[fuel]_{ijt}}}{\left[\dfrac{f_{ijt}}{p_{jt}}\right]} = \mu_j \tag{11}$$

where μ_j denotes the industry markup, while w, r, m and f denote nominal levels (unlike in the previous sections of this paper where they denoted logs) of factor input prices and p is the output price. Dropping industry subscripts and plugging these back into the expression for the production function (totally differentiated), Levinsohn arrives at his final estimating equation:

$$\Delta q_{it} = \mu x_{it} + \Delta\mu(Dx_{it}) + \varepsilon_{it} \tag{12}$$

where D is the trade reform dummy and where

$$x_{it} = \left[\frac{w_{it}}{p_t}\right]\Delta[labor]_{it} + \left[\frac{r_{it}}{p_t}\right]\Delta[capital]_{it} + \left[\frac{m_{it}}{p_t}\right]\Delta[materials]_{it} + \left[\frac{f_{it}}{p_t}\right]\Delta[fuel]_{it} \tag{13}$$

and the changes denote time differences. The markup in Eq. (12) equals μ for the pre-reform period and $\mu + \Delta\mu$ for the post reform period. The parameter of interest is $\Delta\mu$ and indicates the change in the markup following the reform. Levinsohn's primary finding is that $\Delta\mu$ is estimated to be negative for most industries in which tariffs were reduced.

The question then is the following: given Levinsohn's results, how is it that our estimation of the first-order-condition determining the derived demand for labor demand indicates that there were no changes in the relationship between labor demand and wage with changes in openness? To understand the source of the differences between our results and those of Levinsohn's, it is important to note that Levinsohn's estimating equation actually *imposes* that (11) is satisfied with equality across factors — that is to say that in estimating a common markup in (12), it imposes that the first-order conditions (equating the markup to the ratio of price to the marginal cost of producing one extra unit of output by increasing the use of a single input factor) are satisfied across all factors. Levinsohn's stated result from the estimation of (12) that markups declined as a result of greater openness therefore can be read as indicating that the *average* (across all input factors) wedge between marginal products and real factor rewards declined as a result of the trade liberalization in Turkey. Our own re-estimation of (12) allowing for markup coefficients in (11) to differ across factors indicates that this result does not hold factor by factor. Indeed, and this is crucial for our results, the markup reduction that Levinsohn's obtains can be shown to be resulting entirely from changes in the use of material inputs rather than due to changes in labor. Our re-estimation of (12) proceeded in two steps. First, we estimated (12) just as Levinsohn did — in order to verify that we were able to find average (again average across factors) markup reductions just as Levinsohn did. To see if the markup reductions obtain factor by factor, we estimated the following modified equation allowing for the markup coefficient attached to the labor term (μ_L) to be different from the markup coefficient term attached to the remaining factors μ_{-L}:

$$\Delta q_{it} = \mu_{-L}x_{it}^{-L} + \Delta\mu_{-L}(Dx_{it}^{-L}) + \mu_L x_{it}^{L} + \Delta\mu_L(Dx_{it}^{L}) + \varepsilon_{it} \tag{14}$$

where

$$x_{ijt}^{-L} = \left[\frac{r_{ijt}}{p_{jt}}\right]\Delta[capital]_{ijt} + \left[\frac{m_{ijt}}{p_{jt}}\right]\Delta[materials]_{ijt} + \left[\frac{f_{ijt}}{p_{jt}}\right]\Delta[fuel]_{ijt} \tag{15}$$

$$x_{ijt}^{L} = \left[\frac{w_{ijt}}{p_{jt}}\right]\Delta[labor]_{ijt} \tag{16}$$

Table 6 presents our estimates of (12) and (14). Our results match Levinsohn's estimates closely — numerically slightly different although qualitatively the same for each industry.[18] We can clearly see that there was a statistically significant reduction in the markups for almost all industries in which protection declined. As can be clearly seen, the estimates of μ_{-L} and $\Delta\mu_{-L}$ from the estimation of (14) are

Table 6
Reconciliation with Levinsohn's markup results[a]

ISIC code	μ	$\Delta\mu$	μ_{-L}	$\Delta\mu_{-L}$	μ_L	$\Delta\mu_L$
341	0.72 (0.14)	0.66 (0.16)	0.75 (0.15)	0.56 (0.17)	−0.14 (0.63)	3.73 (1.61)
351	2.32 (0.42)	−1.27 (0.48)	2.39 (0.44)	−1.35 (0.49)	−5.06 (12.79)	1.27 (18.57)
352	1.71 (0.35)	−0.85 (0.40)	2.12 (0.32)	−0.90 (0.32)	3.11 (10.61)	−7.10 (14.67)
361	3.57 (0.28)	−2.71 (0.30)	3.40 (0.28)	−2.49 (0.30)	3.39 (1.08)	−6.09 (2.07)
372	0.69 (0.19)	−0.23 (0.22)	−0.79 (0.20)	−0.35 (0.24)	2.30 (1.20)	−1.63 (1.45)
381	0.91 (0.20)	−0.29 (0.20)	1.50 (0.23)	−0.77 (0.22)	−0.16 (0.29)	2.29 (0.58)
382	2.33 (0.14)	−0.98 (0.14)	2.14 (0.13)	−1.14 (0.13)	−1.74 (1.44)	11.98 (1.95)
383	3.28 (0.21)	−1.51 (0.20)	3.43 (0.20)	−1.60 (0.20)	0.29 (1.76)	−0.33 (1.77)
384	0.15 (0.13)	0.55 (0.16)	0.22 (0.13)	0.48 (0.16)	−0.45 (0.28)	2.74 (1.19)
385	1.18 (0.32)	0.34 (0.41)	1.19 (0.33)	0.33 (0.42)	1.44 (1.63)	0.56 (3.36)

[a] Eq. (12):

$$\Delta q_{it} = \mu x_{it} + \Delta\mu(Dx_{it}) + \varepsilon_{it}$$

where

$$x_{ijt} = \left[\frac{w_{ijt}}{p_{jt}}\right]\Delta[labor]_{ijt} + \left[\frac{r_{ijt}}{p_{jt}}\right]\Delta[capital]_{ijt} + \left[\frac{m_{ijt}}{p_{jt}}\right]\Delta[materials]_{ijt} + \left[\frac{f_{ijt}}{p_{jt}}\right]\Delta[fuel]_{ijt}$$

Eq. (14)

$$\Delta q_{it} = \mu_{-L}x_{it}^{-L} + \Delta\mu_{-L}(Dx_{it}^{-L}) + \mu_L x_{it}^{L} + \Delta\mu_L(Dx_{it}^{L}) + \varepsilon_{it}$$

where

$$x_{ijt}^{-L} = \left[\frac{r_{ijt}}{p_{jt}}\right]\Delta[capital]_{ijt} + \left[\frac{m_{ijt}}{p_{jt}}\right]\Delta[materials]_{ijt} + \left[\frac{f_{ijt}}{p_{jt}}\right]\Delta[fuel]_{ijt}$$

and

$$x_{ijt}^{L} = \left[\frac{w_{ijt}}{p_{jt}}\right]\Delta[labor]_{ijt}.$$

[18] We followed the estimation methodology in Levinsohn (1993) closely except that we used the economy-wide, annual interest rate data (borrowing costs) as rental costs, while Levinsohn used a flat rate of 7% throughout.

very close to our estimates of μ and $\Delta\mu$ from (12). However, the estimates of μ_L and $\Delta\mu_L$ are rather different. In other words, the labor part of the x variable had no role to play in the Levinsohn (1993) results. It appears rather that it was the combination of other input factors that generated his results. Thus, it is the average wedge between marginal products and factor rewards that was estimated as the common markup in Levinsohn (1993) and shown to have declined. What we have shown is that this decline did not take place factor by factor.[19] Changes in labor demand in particular do not seem to be playing a role in the drop in average markup results estimated by Levinsohn. This clarifies the apparent lack of inconsistency between our findings and those of Levinsohn (1993).

7. Conclusions

Our analysis of the impact of trade reforms on labor demand elasticities using plant level data over a period spanning major trade reforms in the Turkish economy suggests that the putative linkage between greater trade openness and labor demand elasticities (as suggested by theories of the type we present in this paper) may be empirically quite weak: in the vast majority of the industries we considered separately, we are unable to find statistically and economically significant relationship between these variables. Our results are robust to the type of labor considered (contract labor, overtime labor, female workers, etc.). This non-responsiveness of labor demand elasticity in practice is perhaps explained by a variety of frictions that affect the labor demand decisions of firms.

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[19] We experimented with the division of x in a different way — separating materials from the rest. The coefficients of the real materials changes and those of real materials change times the reform dummy give us roughly the same markup and the markup change. In other words, materials are a major part of the inputs and it seems that it is changes in the use of materials inputs that drive Levinsohn's results.

Table 7
Monte Carlo simulation results

Assumed corr (x,e)	Constant	*Dum*	x	*Dum* $\times x$
0.2	1.03 (0.003)	0.006 (0.003)	−0.79 (0.003)	−0.1 (0.003)
0.4	0.997 (0.003)	0.004 (0.003)	−0.6 (0.003)	−0.1 (0.003)
0.6	1 (0.003)	−0.003 (0.003)	−0.402 (0.003)	−0.101 (0.003)
0.8	1 (0.003)	−0.001 (0.002)	−0.203 (0.002)	−0.098 (0.002)

Appendix A

The Monte Carlo Simulation results presented in Table 7 were derived using the following procedure. With the dependent variable (Y) representing the log demand for labor, and the independent variable X representing the log wage rate, we randomly generate X, a liberalization dummy and the error term e by assuming various correlations (as shown in Table 7) between X and e. The dependent variable is the generated as $Y = 1 - X - 0.1 \times Dum \times X + e$: i.e., assuming that the elasticity before the reforms is -1 and that the change in the elasticity is 0.1 after the reforms. We then regress Y on *Dum*, X and $Dum \times X$. Table 7 presents the results. As expected, our estimates of the coefficient on X are biased — with the bias increasing with larger correlation between X and e. However, as argued in the paper, the coefficient representing the change in elasticity is not biased: the estimates of the coefficient on $Dum \times X$ are not significantly different from -0.1.

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