ELSEVIER Journal of International Economics 39 (1995) 1–25

Journal of
INTERNATIONAL
ECONOMICS

Union wage sensitivity to trade and protection: Theory and evidence

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Received January 1993, revised version received November 1994

Abstract

We develop a model featuring union–firm bargaining, strategic rivalry between the unionized domestic firm and its foreign competitor, and endogenous protection. The model frames a micro-level empirical study of the role of trade and trade policy in union wage determination. The results indicate that (1) trade flows and trade policy influence wages as much as the domestic factors usually considered, (2) imports and tariffs are negatively correlated with wages, and (3) there is little evidence of the trade flows endogeneity suggested by strategic trade theory or the tariff endogeneity that could explain the negative tariff coefficient.

Key words: Wages; Tariffs; NTBs; International trade

JEL classification: F13; F14; J31; J51

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† The authors are grateful to the editorial board and anonymous referees for particularly valuable suggestions. We thank the participants of workshops at L.S.U., McMaster, and the Winter 1993 Econometric Society Meetings. Jim Brander kindly acted as discussant at the 1992 Canadian Economic Association Meetings. We benefited from discussions with Michael Baker, Dwayne Benjamin, S. Lael Brainard, Kevin M. Murphy, Aloysius Siow, Shunichi Tsutsui, and Insan Tunali. Edward Leamer kindly supplied the NTB and tariff data. This research was supported by the Lynde and Harry Bradley Foundation (Gaston) and the Ford Foundation (Trefler).

SSDI 0022-1996(94)01359-4

1. Introduction

In the process of receiving 'fast track' authority for negotiating the North American Free Trade agreement the Bush administration ran head on into one of the largest lobbying campaigns ever mounted by the AFL-CIO (*Wall Street Journal*, 2 April 1991). The decision by the AFL-CIO to make the U.S.–Mexico agreement its top legislative priority marks organized labor as the chief advocate of a more protectionist U.S. trading regime. Organized labor's visibility raises important questions about the role of international trade and protection in union wage determination.

Our investigation of union wages begins with a theoretical framework that builds on advances by Brander and Spencer (1988) and Mezzetti and Dinopoulos (1991). The starting point of these models is a union and firm supplying an internationally traded product in an imperfectly competitive market: the product market is the source of rents that are partially extracted from shareholders by unionized labor. Our focus is on the relationship between observable industry characteristics and the wage negotiated between unions and firms: the industry characteristics include tariffs, non-tariff barriers (NTBs), imports, and exports. The precise model we offer differs from previous work in combining endogenous protection, McDonald and Solow (1981) union preferences, efficient contracting, and simultaneous determination of union wages, domestic output, foreign output, and levels of protection. Other related theoretical contributions include Grossman (1984), Lawrence and Lawrence (1985), Staiger (1988), Brecher and Long (1989), and Dowrick and Spencer (1991). While our theoretical contribution is of independent interest we use it primarily as a guide to empirical work.

Our empirical contribution is to provide a first econometric study of the role of U.S. trade policy in union wage determination. To this end we have constructed a unique, high-quality data set that combines detailed information on labor market outcomes, international trade, and trade protection. The labor market data are from the *Current Population Survey* and include information on the characteristics of over 5,000 workers employed in U.S. manufacturing industries. The protection data are from the 1983 GATT *Tariff Study* and the 1983 UNCTAD *Data Base on Trade Control Measures*. The *Data Base* is the most comprehensive record available of NTBs. Examples of NTBs include quantity restraints (e.g. quotas and VERs), price restraints (e.g. countervailing duties and variable levies), and threats (e.g. anti-dumping investigations).

The paper is closely related to empirical work on union wage determination in an open economy by Lawrence and Lawrence (1985), Katz and Summers (1989), Macpherson and Stewart (1990), Abowd and Lemieux (1991), Freeman and Katz (1991), and Fung and Huizinga (1991). We depart from previous research in combining the following three features for the

United States: (i) the use of micro-level wage data rather than industry-level wage data; (ii) a focus on union members rather than all workers; and (iii) data not only on trade flows, but also on tariffs and NTBs which are direct policy variables.

As with any study using a new approach, here combining micro-level wage data standard in the labor literature with industry-level protection data standard in the international trade literature, the results raise as many questions as they answer. Chief among these is the negative coefficient on tariffs in the wage equation. Although this is shown to be consistent with the theory, it is nevertheless surprising. As shown in Subsection 4.3, it apparently cannot be explained by the endogeneity of protection (low wages lead to high tariffs) or by the endogeneity of trade flows (import and export levels are strategic choice variables of firms).

A form of the negative tariff coefficient appears in all three previous cross-industry micro-level wage studies examining tariffs. It appears for non-union Canadian workers (Fung and Huizinga, 1991) and for all U.S. workers, union and non-union pooled (Gaston and Trefler, 1993, 1994b). Gaston and Trefler (1993) was primarily concerned with NTBs: we concluded that for a variety of definitions of NTBs, the impact of NTBs was sensitive to the choice of specification and always statistically insignificant. Gaston and Trefler (1994b) demonstrated that the negative tariff coefficient was robust to a variety of alternative statistical assumptions regarding one- versus two-stage estimators, non-spherical errors, outliers, regressor exogeneity, choice of regressors, etc. The current paper demonstrates that the negative tariff effect is strictly a union phenomenon by showing that there is no relationship between tariffs and non-union wages. More generally, we show that there is a fundamental difference between union and non-union wage responses to trade and trade policy. Furthermore, while our earlier papers were empirical explorations devoid of economic modelling, our current paper develops a detailed theory capable of explaining the negative tariff result and other union–non-union differences.

A different approach to estimating the effect of trade policy on union wages uses time-series data rather than the cross-section data of this study. Gaston and Trefler (1994a) examined the impact on wages of the 1989–91 bilateral tariff reductions negotiated under the Canada–U.S Free Trade Agreement. In a different vein, Grossman (1987) and Revenga (1992) exploited time-series data to estimate the effect of unit-value import prices on average industry wages. All three studies estimated small wage elasticities. To investigate further, Revenga instrumented import prices using industry-specific measures of trading partner exchange rates and foreign production costs. This reflected her interest in the effect of exchange rate movements rather than tariffs and NTBs. Subsection 4.4 further examines time-series issues.

2. The role of trade and protection in wage determination

Our analysis of the role of trade and protection in wage determination is set against the backdrop of wage–employment bargaining between a union and management. On the labor market side, the model is a variant of McDonald and Solow's (1981) efficient bargaining with von Neumann–Morgenstern union preferences. On the product market side, two different specifications of international trade are used. In Subsection 2.1, the basis of trade is comparative advantage: import competition and export performance are treated as *exogenous* indicators of domestic competitiveness. In Subsection 2.2, the basis of trade is strategic rivalry between domestic and foreign firms, as in Brander and Spencer (1985): imports and exports are treated *endogenously*. In Subsection 2.3, levels of protection are modelled endogenously.

2.1. Imports and exports as indicators of comparative advantage

Consider a partial equilibrium model in which there are no strategic interactions between firms in the industry. By assumption, the market structure generates rents that are shared between firms and unions. Attention is focused on a representative firm facing a union. The firm's total revenue is $R(n, \rho)$, where n is employment and ρ is a shift variable indicating the competitive position of the firm. Higher values of ρ are associated with greater total revenue, i.e. $R_\rho > 0$. Assume $R_n > 0$ and $R_{nn} < 0$.

The union and the firm bargain over wage–employment (w, n) contracts. We adopt the McDonald and Solow (1981) specification of union preferences. Without loss of generality normalize n so that it denotes the probability of employment. Let ω be the alternative wage of a union member who, with probability $1 - n$, fails to gain employment under the union contract. Union preferences over (w, n) lotteries are given by $nU(w) + (1 - n)U(\omega)$. We assume risk aversion.

Bargaining over (w, n) contracts is assumed efficient and occurs simultaneously over wages and employment. The choice from the set of efficient contracts is the one that maximizes the generalized Nash product:

$$G(w, n) = (R(n, \rho) - wn - \pi)^{1-\alpha}(nU(w) + (1 - n)U(\omega) - U(\omega))^{\alpha},$$

where $(\pi, U(\omega))$ is the disagreement point and $\alpha \in (0, 1)$ indexes collective-bargaining strength. We assume the Nash solution lies in the interior of the choice set ($n \in (0, 1)$ and $w > \omega$) and that G is strictly concave so that the solution is unique and may be characterized by first-order conditions.

The first-order conditions determining the efficient contract are:

$$\frac{\partial \ln(G)}{\partial w} = -\frac{(1-\alpha)n}{R(n,\rho) - wn - \pi} + \frac{\alpha U_w(w)}{U(w) - U(\omega)} = 0\,, \tag{1}$$

$$\frac{\partial \ln(G)}{\partial n} = \frac{(1-\alpha)(R_n(n,\rho) - w)}{R(n,\rho) - wn - \pi} + \frac{\alpha}{n} = 0\,. \tag{2}$$

Substituting (1) into (2), yields the familiar set of efficient contracts:

$$\frac{U(w) - U(\omega)}{U_w(w)} = w - R_n(n,\rho)\,. \tag{3}$$

Our focus is on industry characteristics that affect the negotiated wage $w = w(\alpha, \pi, \omega, \rho)$. Proposition 1 collects the relevant results. All proofs appear in the Mathematical appendix as do results about $n = n(\alpha, \pi, \omega, \rho)$.

Proposition 1 (Comparative advantage). $w = w(\alpha, \pi, \omega, \rho)$, $\partial w/\partial\alpha > 0$, $\partial w/\partial\pi < 0$, $\partial w/\partial\omega > 0$, *and* $\partial w/\partial\rho$ *has an indeterminate sign.*

The proposition states that the contracted wage is higher the greater is union bargaining strength, the worse are the firm's outside profit opportunities, and the better is the alternative wage. It also states that a higher level of demand (ρ) need not lead to a higher union wage; in particular, higher levels of protection or lower levels of import competition need not translate into a higher wage. However, $\partial w/\partial\rho < 0$ implies $\partial n/\partial\rho > 0$. That is, if trade protection leads the union to negotiate wage cuts, then in return the union receives implicit or explicit employment guarantees. A negative relationship between protection and wages appears in Lawrence and Lawrence (1985) and Staiger (1988).

In terms of empirical implementation, ρ will be measured by imports, exports, tariffs, and NTBs. This raises the question of whether, after controlling for imports, levels of protection have any effect on R and hence w. Suppose the domestic firm is a monopolist using focal point pricing, i.e. pricing just below the world price plus tariff. A rise in the tariff does not affect imports since imports are excluded, but permits the monopolist to raise its price. Thus, the tariff increases R without a corresponding change in imports. An example of focal point pricing is the American Selling Price NTB in the chemical industry. This is a simple example because it involves prohibitive protection; however, a richer model of firm interaction would allow for a richer role of protection independent of imports.

For our empirical purposes, the assumption of efficient contracting is easily replaced by a monopoly union assumption where the union chooses w and the firm responds by choosing n such that (w, n) lies on the firm's labor demand schedule, R_n (e.g. Grossman, 1984; Lawrence and Lawrence, 1985). Then the (w, n) contract satisfies

$$\max_{w} nU(w) + (1-n)U(\omega) \quad \text{s.t.} \quad R_n(n, \rho) = w.$$

The optimal wage depends on ω and ρ only, $\partial w/\partial\omega > 0$, and $\partial w/\partial\rho$ has an indeterminate sign that depends among other things on the third derivative, $R_{nn\rho}$. Thus, assuming a monopoly union in place of efficient contracting does not affect the conclusions about the relation of w to the outside option ω or to the international trade variables ρ. However, in a regression setting it imposes exclusion restrictions on regressors measuring α and π. These restrictions can be examined empirically.

2.2. International trade with strategic rivalry

We now explicitly introduce a foreign firm that is a strategic rival to the domestic firm. There are two firms producing for the domestic market, one domestic firm with output x and one foreign firm with output y. Attention is restricted to tariffs since the case of a binding quota is straightforward: if the quota binds at a level $y = \bar{y}$, then the analysis is the same as in the previous section with imports set exogenously at $\bar{y}$. The foreign firm has profits $R^*(y, x) - w^*y - \tau y$ from operating in the domestic market, where R^* is the foreign revenue function, w^* is the exogenous foreign wage, and τ is the domestic tariff which the foreign firm must pay. The domestic firm has profits $R(x, y) - wx$, where $R(x, y)$ is the revenue function and w is the wage. We continue to denote the domestic revenue function by R even though it is now a function of (x, y) rather than (n, ρ) as in the previous section. Assume $R_x > 0$, $R_{xx} < 0$, $R_y < 0$, $R^*_y > 0$, $R^*_{yy} < 0$, and $R^*_x < 0$. In line with Brander and Spencer (1988) and Mezzetti and Dinopoulos (1991) assume that one unit of labor is needed to produce one unit of output so that $n = x$. Sufficient conditions for stability are set out in the Mathematical appendix. Finally, we assume $R_{xx}R^*_{yy} - R_{xy}R^*_{yx} > 0$, which is the usual stability condition for the game between the domestic and foreign firms when wages are set exogenously.

The domestic firm bargains with the risk-averse union over wage–output (w, x) contracts. Competition between the domestic and foreign firms is treated as a Cournot quantity (x, y) game. w, x, and y are simultaneously determined. This combines Mezzetti and Dinopoulos's (1991) staging with Brander and Spencer's (1988) preferences. The first-order conditions determining w and x are as in Eqs. (1) and (2) with $R(n, \rho)$ replaced by $R(x, y)$ and n by x. In addition, for the foreign firm we have the first-order condition:

$$R^*_y - w^* - \tau = 0\,. \tag{4}$$

The relevant comparative statics are summarized by the next proposition.

Proposition 2 (Strategic rivalry). $w = w(\alpha, \pi, \omega, \tau)$, $\partial w/\partial \alpha > 0$, $\partial w/\partial \pi < 0$, *and both* $\partial w/\partial \omega$ *and* $\partial w/\partial \tau$ *have indeterminate signs.*

The proposition states that higher wages are associated with greater union bargaining strength (α) and lower outside options for the firm (π). The effect of a tariff on wages and output is ambiguous even under the strategic substitutes assumption, a point discussed in Brander and Spencer (1988).[1] The strategic trade model just outlined is intended to illustrate the endogeneity of trade flows. Other strategic trade models could have been used to make this point, but we do not present them because we will not attempt to test or discriminate between alternative strategic trade models.

2.3. Endogenous protection and trade

Falling wages and declining employment figure prominently in Commerce Department, ITC, and Presidential decisions to protect an industry. The responsiveness of trade policy to wages and employment may be captured by the function $\tau = S^{\tau}(w, x)$, where $S^{\tau}_{w} < 0$ and $S^{\tau}_{x} < 0$. (Only $S^{\tau}_{x} < 0$ will be needed.) Modelling endogenous tariffs is complicated by the many cases in which foreign firms have behaved strategically in choosing a level of exports to the United States that is low enough to prevent retaliation by U.S. policy-makers. We model this as a non-cooperative Cournot–Nash game between producers and policy-makers. $\tau = S^{\tau}(w, x)$ is the Cournot–Nash reaction function of policy-makers. Producers have reaction functions $w = S^{w}(\tau)$, $x = S^{x}(\tau)$, and $y = S^{y}(\tau)$, where S^{w}, S^{x}, and S^{y} are the equilibrium values of w, x, and y, respectively, from the previous section.

We retain the same assumptions made in the previous section. As well, we assume that the game between producers and policy-makers is stable. For the first time we will also need that x is a strategic substitute for y ($R_{xy} < 0$).

Proposition 3 (Endogenous trade and protection). Assume $R_{xy} < 0$. *Then* $w = w(\alpha, \pi, \omega)$, $\partial w/\partial \alpha > 0$, $\partial w/\partial \pi < 0$, *and* $\partial w/\partial \omega$ *has an indeterminate sign.*

The results do not differ from Proposition 2 except in the magnitude of the derivatives and the absence of τ as an argument in the wage function. The basic insight is that the building block union–firm bargaining that de-

[1] In this model τ enters w only through y, which means that tariffs do not play a role independent of imports. However, several modifications of the model provide an independent role for tariffs. For example, suppose there are fixed costs and the home firm moves first by choosing a limit output that excludes the foreign firm from the domestic market. Since the limit output will under weak conditions be a decreasing function of τ, higher values of τ lead to higher values of R even though y is unchanged.

termines how α, π, and ω affect w is unaltered by endogenous protection. Of course, the co-movement of w with τ, x, and y, which is the focus of our empirical work, is now considerably more complicated.

3. Methodology

An important innovation of this paper is that it combines detailed data on trade protection with detailed micro data on individuals' labor characteristics and earnings. All data apply to U.S. manufacturing in 1983 except where noted. We know of no U.S. study relating union wages to trade protection. The trade protection data are from the GATT *Tariff Study* and the extensive UNCTAD *Data Base on Trade Control Measures*. NTBs are measured as the proportion of imports covered by an NTB or what is termed a coverage ratio. Tariffs are average tariffs and are aggregated using import weights. Trade data are from unpublished Bureau of Commerce sources. Imports and exports are scaled by domestic output plus net imports. We merge the industry-level trade and trade protection measures with micro data taken from the 1984 *Current Population Survey* (CPS). The Data appendix describes how our sample of 5,474 union workers was selected. We obtained the same results using 1982 and 1983 CPS data, but chose 1984 in order to fully avoid the 1982 recession.

To determine the role of industry characteristics such as trade and protection in wage determination we use the interindustry wage differentials approach (e.g. Dickens and Katz, 1987; Krueger and Summers, 1988). An individual's wage depends at least as much on his or her own human capital, demographic, geographic, and occupational characteristics as it does on the characteristics of the individual's industry of affiliation. Thus, we begin by postulating an earnings equation that depends on the individual's characteristics as well as his or her industry of affiliation:

$$\log(W_{ij}) = H_i\beta_H + D_jW_j^* + \varepsilon_{ij}, \quad i = 1, \ldots, I, \quad j = 1, \ldots, J , \tag{5}$$

where W_{ij} is the wage of individual i in industry j, H_i is a vector of characteristics of individual i, D_j is a vector of mutually exclusive dummy variables indicating industry of affiliation, ε_{ij} is an error term, and β_H and W_j^* are parameter vectors. W_j^* is the interindustry wage differential or wage premium for industry j: it is that portion of an individual's wage which cannot be explained by his or her observable individual characteristics H_i, but which is explained by his or her industry of affiliation. The extent to which wage premia reflect unobservables or non-competitive rents is an open question.

The H_i and β_H are not the main focus so we comment on them only briefly. The H_i characteristics we chose form a conventionally selected set of included variables and the coefficient signs and magnitudes we estimated are very similar to those reported by many previous researchers.[2]

It is tempting to include industry-level characteristics such as trade and protection directly into the earnings equation:

$$\log(W_{ij}) = H_i\beta'_H + P_j\beta'_P + X_j\beta'_X + \varepsilon'_{ij}, \quad i = 1, \ldots, I, \quad j = 1, \ldots, J, \tag{6}$$

where P_j is a vector of international trade and protection factors measuring ρ and τ and X_j is a vector of domestic factors measuring α, π, and ω. Unfortunately, OLS estimation of Eq. (6) may lead to standard errors that exaggerate the significance of the included industry-level variables (Moulton, 1986). The problem is that there are typically industry-level error components so that $\varepsilon'_{ij} = \lambda_j + \mu_{ij}$, where λ_j is the 'fixed effect' or disturbance common to all individuals in industry j. For example, λ_j may reflect selectivity issues associated with unobserved worker heterogeneity. One approach to estimating error components is GLS. Another common approach (see citations in Dickens and Katz, 1987, p. 65) proceeds in two stages: (i) use OLS to estimate the W_j^* in Eq. (5) and (ii) use OLS to estimate

$$W_j^* = P_j\beta_P + X_j\beta_X + \nu_j, \quad j = 1, \ldots, J. \tag{7}$$

This two-stage procedure has the advantage that it is easy to use, allowing us to examine a wide variety of specifications even in the more complicated instrumental variables setting used below. In Subsection 4.3 we address the more difficult problem that occurs when the fixed effects λ_j are correlated with the exogenous regressors.

4. Empirical results

4.1. Data preview

The sample consists of 5,474 union workers of which 77% are males, 72% have 12 or less years of education, and 50% have 22 or less years of labor force experience. OLS estimation of Eq. (5) yields estimates of the wage

[2] The coefficients (standard errors) are: Schooling 0.031 (0.0020); labor force experience 0.013 (0.0013); experience squared −0.00019 (0.000025); employed full-time 0.058 (0.020); male 0.16 (0.013); white 0.064 (0.012); household head 0.030 (0.011); married 0.028 (0.011); lives in South −0.054 (0.010); lives in inner city −0.025 (0.0093); veteran 0.030 (0.0095); engineer-scientist 0.31 (0.040); white collar 0.098 (0.018); skilled 0.16 (0.015); and semi-skilled 0.058 (0.013), where the sample consists of 5,474 union workers and $R^2 = 0.42$.

premia W_j^*. Table 1 previews the largest and smallest wage premia. The wage premia are normalized so that their employment-weighted mean is zero. For example, after controlling for observable human capital characteristics apparel workers earn 39% less than the average manufacturing worker. The wage premia are similar to the wage premia for all workers (union and non-union) reported in Krueger and Summers (1988, Table A1). Comparing the rows of averages for high and low wage premia industries, a simple characterization appears: low wage premia industries are high-tariff,

Table 1
Data preview

Industry	Smallest wage premia					
	Wage premium	Standard error[b]	Tariffs	NTBs	Exports	Imports
Wood products[a] (241)	−40	7.1	6	4	3	9
Apparel[a] (152)	−39	7.7	12	48	3	5
Footwear[a] (221)	−38	6.0	9	92	1	43
Apparel[a] (151)	−33	3.9	23	12	1	19
Textile knitting mills (132)	−31	7.9	22	40	1	.0
Pottery (261)	−27	9.3	12	3	8	36
Misc. manufacturing (391)	−26	5.4	5	1	6	22
Textiles[a] (150)	−25	10.6	7	12	8	11
Textiles[a] (140)	−23	5.1	14	57	3	6
Cutlery, tools, hardware (281)	−17	5.9	9	0	6	10
Unweighted average	−30	6.9	12	27	4	16
Industry	**Largest wage premia**					
	Wage premium	Standard error[b]	Tariffs	NTBs	Exports	Imports
Pulp, paper & paperboard (160)	12	4.3	0	0	8	13
Aircraft & parts (352)	13	4.4	0	12	31	6
Motor vehicles & parts (351)	13	4.0	2	37	8	22
Beverages (120)	14	5.0	5	95	1	7
Engines & turbines (310)	14	5.4	3	15	28	14
Tobacco (130)	15	7.0	11	14	8	3
Tires & inner tubes (210)	15	5.5	4	30	3	13
Logging (230)	18	7.3	0	15	15	1
Newspaper publish-print (171)	20	5.2	0	0	0	0
Petroleum refining (200)	24	5.3	0	0	3	9
Unweighted average	16	5.3	3	22	11	9

Note: All figures are multiplied by 100 in order to express them as percentages.
[a] Data relate to only that part of the industry indicated by the parenthesized Census Industrial Classification codes.
[b] These are OLS standard errors of W_j^* from Eq. (5). The employment-weighted mean of the wage premia is zero.

import-competing industries while high wage premia industries are low-tariff, export-oriented industries.

4.2. Exogenous trade and protection

In order to assess the role of trade and protection in wage determination we estimate Eq. (7) while treating trade and protection exogenously. P_j is taken to be tariffs, NTBs, imports, and exports. The reasons for including both imports and protection were outlined above. An additional econometric reason is that in a cross-industry regression setting imports and exports may not completely control for interindustry differences in revenue functions. For example, suppose that two industries have differing demand elasticities, but identical import levels because of tariff protection. Since the industry with the less-elastic demand will have the higher tariff ceteris paribus, the tariff level conveys information about R and hence W_j^* in the two industries even though import levels are identical.

X_j in Eq. (7) is taken to be measures of α, π, and ω. Also included is a proxy for the size of labor market rents, the *four-firm concentration ratio*. A firm's bargaining strength, $1-\alpha$, is related to the firm's ability to send production abroad where it is out of the union's reach. (Mezzetti and Dinopoulos, 1991, incorporate offshore sourcing into π and ω.) While comprehensive data on offshore sourcing is not available one proxy is *MOFA imports*, i.e. goods entering the United States that were purchased by U.S. multinationals from their majority-owned foreign affiliates (MOFAs). A defect of this proxy is that at least some MOFA imports are complements to rather than substitutes for the output produced by the U.S. parent. Also, MOFA imports are a component of total imports and so will be subject to concerns addressed below about regressor endogeneity. An alternative proxy that is correlated with MOFA imports, but more likely to be exogenous, is *MOFA employment*. MOFA imports and employment are scaled by the output of the U.S. parent. For π, if no agreement is reached a firm loses the rents associated with firm-specific investments in physical capital (measured by the *capital/labor ratio*) and intangible capital (measured by the *profit rate*[3]). Letting r be rents, in the event of agreement the firm earns $R - wn + r$. Comparing this with $R - wn - \pi$ in Eqs. (1) and (2), all results for π apply equally for $-r$. Thus, π is negatively related to the capital/labor ratio and the profit rate.

[3] The profit rate is the cumulative rate of return over the year, averaged over the 1982–84 period. Profit rates are calculated for (publicly traded) stocks on the CRSP database which are listed on either the NYSE or AMEX stock exchanges. We are indebted to Marlene Puffer for supplying the data.

The empirical work on union–firm bargaining has been preoccupied with the measure of the alternative wage ω because key results typically depend on the measure used. For example, Brown and Ashenfelter (1986) used 11 different measures. What follows is a list of the measures we will use together with a description of their strengths and weaknesses. The cleanest approach is to exclude any measure of ω from the regressions since ω is controlled for in the first-stage regressions by the schooling, experience, occupation, and other human capital controls. Another approach introduces explicit measures of ω. The *offshore wage* is a measure of the earnings of foreign workers employed by MOFAs and reflects technologically determined differences in average skill levels across industries. The measure suffers from being indirect; however, unlike the other proxies to be discussed it is a pure industry variable that is not defined using the CPS micro data from which the wage premia were constructed. The *union–non-union wage differential* is calculated by pooling union and non-union workers and estimating a variant of Eq. (5) that includes both industry dummies and an interaction of union status with the industry dummies. If the union–non-union differential is large, then possibly the difference between the union wage premium and ω is large. A drawback of this variable is that, almost by construction, a high level of the union wage premium is associated with a high level of the union–non-union differential, thus raising questions about exogeneity. A worker's alternative wage depends on his human capital characteristics of which *schooling* is arguably the most important component. However, schooling in wage premia equations can be interpreted in a variety of ways, not just as ω. Furthermore, the rationale for including schooling leads one to include many other human capital characteristics such as the proportion of workers in the industry by occupation. A measure of ω that combines features of the union–non-union differential and schooling is the *wage decile*. It is the tenth percentile of the industry wage after controlling for the occupational mix of the industry. Specifically, let $W_{[10],jk}$ be the tenth percentile of the wage distribution of occupation k in industry j. The wage decile is defined as $\Sigma_k W_{[10],jk} p_{jk}$, where p_{jk} is the proportion of workers in occupation k in industry j. The wage decile is related to a measure used by Abowd and Kramarz (1993) and is our a priori preferred measure.[4] This completes the discussion of the regressors used in what follows.

[4] See the Data appendix for further discussion of data construction. The unweighted sample means (standard deviations) are: union wage premia −0.0313 (0.146); non-union wage premia 0.00259 (0.0975); tariffs 0.0528 (0.0454); NTBs 0.172 (0.238); imports 0.0915 (0.101); exports 0.0793 (0.0823); four-firm concentration ratio 0.360 (0.148); profit rate 0.264 (0.0978); capital/labor ratio 0.0514 (0.0629); union–non-union differential 0.000549 (0.107); offshore wage 7.866 (2.445); wage decile 1.578 (0.203); schooling 12.247 (0.740); MOFA imports 0.00936 (0.0103); and MOFA employment 0.00307 (0.00176).

Column (1) of Table 2 reports the results when ω is proxied by the wage decile. Column (2) reports the minimum and maximum coefficient estimates across 10 specifications: ω treated in five different ways (excluded or measured by one of four proxies) and $1-\alpha$ treated in two different ways (MOFA imports or MOFA employment). Across the 10 specifications each domestic regressor either has the sign predicted by Proposition 1 or the sign

Table 2
Wage premia regressions

Dependent variable	Union wage premia (W_i^*)					Union wages ($\log W_{ij}$)	Non-union wage premia (W_i^*)
	(1)	(2)		(3)	(4)	(5)	(6)
P: International		min	max				
Tariffs	−1.342	−1.517,	−1.032	−1.760	·	−1.064	−0.094
	(0.320)			(0.334)		(0.097)	(0.165)
NTBs	0.044	−0.013,	0.074	0.038	·	0.024	−0.011
	(0.055)			(0.063)		(0.020)	(0.029)
Imports	−0.348	−0.480,	−0.269	−0.370	·	−0.374	0.040
	(0.138)			(0.142)		(0.061)	(0.071)
Exports	−0.049	−0.131,	0.234	0.168	·	−0.178	0.157
	(0.176)			(0.182)		(0.063)	(0.091)
X: Domestic							
Four-firm concentration	0.131	0.120,	0.270	·	0.037	0.182	−0.026
	(0.105)				(0.107)	(0.032)	(0.054)
Profit rate	0.144	−0.095,	0.144	·	0.121	0.078	−0.102
	(0.156)				(0.165)	(0.051)	(0.081)
Capital/labor ratio	0.325	0.304,	0.699	·	0.318	0.350	0.244
	(0.228)				(0.265)	(0.084)	(0.117)
Wage decile	0.264	0.264,	0.271	·	0.412	0.248	0.369
	(0.085)				(0.091)	(0.035)	(0.044)
MOFA employment	−1.069	−1.069,	7.453	·	−9.264	−3.419	4.983
	(7.614)				(8.588)	(2.354)	(3.920)
Intercept	0.447	−0.811,	−0.027	0.076	−0.714	0.848	−0.635
	(0.140)			(0.031)	(0.148)	(0.065)	(0.072)
MOFA imports	·	−1.959,	0.089	·	·	·	·
Offshore wage	·	0.002,	0.002				
Union–non-union	·	0.656,	0.668	·	·	·	·
Schooling	·	0.061,	0.066	·	·	·	·
n	68	68		68	68	5,474	68
R^2	0.595			0.412	0.410	0.387	0.759

Note: Standard errors are in parentheses. Columns (1)–(4) are alternative OLS estimates of Eq. (7). Specifically, column (2) reports the minimum and maximum coefficient bounds across the 10 specifications discussed in the text. Column (5) is the OLS estimate of Eq. (6) (the human capital coefficients β_H' are not reported). Column (6) is the OLS estimate of Eq. (7) using non-union wage premia.

is not robust, i.e. it is sensitive to the choice of specification. The measures of ω always have the predicted sign and are statistically significant. The four-firm concentration ratio and the capital/labor ratio coefficients always have the correct sign and, except when ω is treated as the wage decile, are statistically significant. These results together with the reasonable R^2 goodness-of-fit measure at least partly support our use of the McDonald–Solow, Brander–Spencer, and Mezzetti–Dinopoulos framework for analyzing the role of trade and protection in wage determination.

For the international regressors, greater exposure to international trade (imports and exports) leads to lower wages. Across the 10 specifications the import coefficient is robust and always statistically significant. The export coefficient is neither robust nor ever statistically significant. Using wage premia calculated by pooling union and non-union workers, Katz and Summers (1989) found a positive bivariate correlation between wage premia and exports. They used this finding to make a (limited) case for an active, export-oriented trade policy. Our multivariate results vitiate their policy conclusions for heavily unionized industries. Higher levels of tariffs lead to lower wages, an effect that is always statistically significant. Higher levels of NTBs lead to higher wages, but the coefficient is neither robust nor statistically significant. The result for NTBs does not change when NTBs are disaggregated by type, a point explored in detail in Gaston and Trefler (1993) for the case where union and non-union workers are pooled. Using a similar methodology Fung and Huizinga (1991) estimated a statistically significant positive effect of NTBs on the wages of Canadian workers. Our results about coefficient signs, robustness, and statistical significance will appear repeatedly in this paper so that it will simplify exposition to discuss a single specification, that reported in column (1).

A feature of Table 2 revealed by the R^2 values in columns (1), (3), and (4) is the performance of the domestic regressors relative to the international regressors. The domestic regressors explain between 18.3% (0.595 − 0.412) and 41.0% of the cross-industry variation in union wage premia. Likewise, the international regressors explain between 18.5% and 41.2% of the variation. Furthermore, the regressors with by far the largest beta coefficients are imports (−0.24), the wage decile (0.37), and tariffs (−0.42). Thus, international trade and protection appear as important determinants of the cross-industry variation in union wage premia.

We also considered three types of heteroskedasticity corrections and weighting schemes: (1) a White correction (the White test rejects heteroskedasticity); (2) scaling by $n_j^{1/2}$, where n_j is the number of workers in industry j; and (3) scaling by the standard errors of the estimated wage premia. For each of these there was a reduction in standard errors and, in cases (2) and (3), little change in coefficient estimates.

Column (5) of Table 2 reports OLS estimates of the one-stage Eq. (6)

procedure. As expected, it yields similar coefficient estimates and inappropriately small standard errors. Thus, the two-stage procedure does not introduce spurious coefficient estimates.

The good fit and reasonableness of most of the parameter estimates in Table 2 lends support to the model outlined in Section 2. However, one can imagine other theories that imply similar estimating equations without any appeal to the role of unions. Efficiency wage theories are a prominent example (e.g. Weiss, 1980). In order to disentangle union–firm bargaining theory from these other theories we present results for the obvious control group, namely the non-union sample. In the first stage [Eq.(5)] we estimated wage premia for 15,223 non-union workers. In the second stage [Eq. (7)] we regressed non-union wage premia on the same set of regressors as considered for the union sample. The second-stage results appear in column (6) of Table 2. The difference between the union and non-union international coefficients [columns (1) and (6)] is statistically significant at the 1% level with $F_{4,2(68-10)} = 5.13$. This is particularly obvious for the tariff coefficient which is small and statistically insignificant for the non-union sample. Indeed, when ω is treated as the wage decile or schooling, the international regressors as a group are not statistically significant in the non-union regressions. Thus, there is a fundamental difference between union and non-union wage responses to trade and trade policy. One explanation for the asymmetry is that unions face a wage–employment trade-off not available to non-union workers. In response to tariffs, union workers may negotiate a low-wage contract in return for implicit or explicit guarantees of higher employment levels. In contrast, non-union workers do not have this option. If the wage is set competitively, then it is fixed at the level of the exogenous outside option ω and so is independent of the tariff level. Likewise, if the wage is set at its efficiency wage level as in Weiss (1980), then the firm never reduces wages in return for higher employment since such a strategy attracts low-quality workers. These observations are consistent with our empirical findings.

There is increasing evidence that it can be misleading to group all workers together as we have done: in the 1980s very different wage trends emerged for males relative to females, for college graduates relative to less-educated workers, and for experienced workers relative to inexperienced workers (Bound and Johnson, 1992). To investigate, we re-estimated wage premia [Eq. (5)] separately for different groups. Because of small sample sizes, some industries were omitted for some groups. We then estimated Eq. (7) separately for each group. The results appear in Table 3. In order to focus on differences between groups, all the regressions in Table 3 use the largest common set of industries for which wage premia are available.

Reading across different unionized groups (columns of Table 3), the male, less-educated, and experienced samples each behaves the same as the

Table 3
Wage premia regressions: By worker characteristic

	Union sample				Non-union sample			
Dependent variable: wage premia (W_i^*)	All	Sex Male	Schooling $S \leq 12$	Experience $E \geq 22$	All	Sex Male	Schooling $S \leq 12$	Experience $E \geq 22$
P: International								
Tariffs	−1.071	−0.953	−1.012	−0.877	−0.018	0.146	−0.124	0.033
	(0.356)	(0.388)	(0.365)	(0.377)	(0.183)	(0.200)	(0.216)	(0.228)
NTBs	0.045	0.037	0.043	0.048	0.001	−0.014	−0.008	−0.040
	(0.061)	(0.066)	(0.062)	(0.064)	(0.031)	(0.034)	(0.037)	(0.039)
Imports	−0.540	−0.412	−0.510	−0.576	−0.074	−0.039	−0.035	−0.020
	(0.181)	(0.197)	(0.186)	(0.192)	(0.093)	(0.102)	(0.109)	(0.116)
Exports	−0.082	−0.024	−0.172	−0.047	0.189	0.220	0.092	−0.049
	(0.181)	(0.197)	(0.185)	(0.192)	(0.093)	(0.102)	(0.109)	(0.116)
X: Domestic								
Concentration	0.101	0.015	0.145	0.112	0.012	0.000	0.031	0.024
	(0.112)	(0.123)	(0.115)	(0.119)	(0.058)	(0.063)	(0.068)	(0.072)
Profit rate	0.161	0.243	0.118	0.123	0.093	0.062	0.062	−0.010
	(0.164)	(0.178)	(0.168)	(0.173)	(0.084)	(0.020)	(0.099)	(0.105)
Capital/labor ratio	0.291	0.276	0.346	0.217	0.262	0.204	0.169	0.071
	(0.233)	(0.254)	(0.239)	(0.247)	(0.120)	(0.131)	(0.141)	(0.150)
MOFA employment	−3.259	−5.971	−1.667	−6.769	3.399	2.295	3.826	−1.717
	(7.784)	(8.482)	(7.987)	(8.247)	(3.995)	(4.384)	(4.714)	(4.995)
Wage decile	0.274	0.298	0.258	0.257	0.317	0.292	0.397	0.305
	(0.098)	(0.106)	(0.100)	(0.103)	(0.050)	(0.050)	(0.059)	(0.063)
Intercept	−0.437	−0.498	−0.415	−0.393	−0.554	−0.518	−0.642	−0.476
	(0.148)	(0.162)	(0.152)	(0.157)	(0.076)	(0.083)	(0.090)	(0.095)
$\bar{R}^2$	0.497	0.399	0.467	0.434	0.693	0.603	0.684	0.476

Note: Standard errors are in parentheses. Each column reports OLS estimates of Eq. (7). The dependent variable (wage premia) was calculated by estimating Eq. (5) only for those workers indicated in the column heading. The independent variables are the same for each regression (column). Each regression uses the same 56 industries (observations).

pooled ('All') sample. Thus, aggregation is warranted. Note that the tariff coefficient for all union workers is − 1.071 in Table 3 compared with − 1.342 in Table 2. The difference is entirely due to the choice of industries in the second-stage [Eq. (7)] regression: the 12 omitted industries in Table 3 almost all have above-average levels of tariffs. For the non-union sample, male and experienced workers have positive tariff coefficients. Furthermore, the export coefficient is statistically positive for male non-union workers. Thus, union–non-union differences in wage sensitivity to trade and protection is especially pronounced for certain groups such as males.

4.3. Endogenous trade and protection

In this subsection we examine the endogeneity of trade and protection suggested by our theoretical models of Subsections 2.2 and 2.3. Our strategy is to instrument tariffs, NTBs, imports, and exports along the lines of Trefler (1993). Many researchers have emphasized that trade models with imperfect competition complement factor endowments trade models, e.g. Helpman and Krugman (1985, p. 57). Factor endowments models predict that imports depend on factor shares. Factor shares are the sum of 'direct' plus 'indirect' (in an input–output sense) factor shares. For example, in order for industry j to export one dollar of goods it must directly produce one dollar of goods. The associated factor payments are the direct factor shares. Furthermore, the industry must purchase intermediate goods from other industries and these other industries must in turn purchase intermediate goods from yet other industries. The factor payments generated by this multiplier process are the indirect factor shares. Factor shares, direct plus indirect, capture general equilibrium interactions between all industries in the economy and are accordingly constructed from data on all industries. This lessens the likelihood that a factor share in industry j is correlated with the disturbance in industry j. We use the 13 factor shares described in Trefler (1993) as well as the exogenous (domestic or X_j) regressors in the wage premium equation.

Columns (1) and (2) of Table 4 report the OLS and instrumental variables (IV) estimates. The Hausman test leads one to weakly accept the hypothesis of endogeneity at the 6.1% level. This is consistent with Fung and Huizinga's (1991) rejection of tariff endogeneity and acceptance of NTB endogeneity for Canada. There is nothing unusual about the first-stage regressions. Furthermore, the instruments pass the Basmann overidentification test and, correspondingly, the conclusions do not change much when subsets of the 13 factor share instruments are used. The only notable difference between the OLS and IV results is that the tariff coefficient becomes more negative. Possibly this reflects the poor properties of IV estimators when there is a small sample and four endogenous regressors. To address this we examine a model with only two endogenous regressors, tariffs and imports, and with NTBs and exports omitted from the regression. As the OLS results of columns (1) and (3) indicate, omitting NTBs and exports has little impact on the remaining coefficients. It does, however, produce a clearer acceptance of the endogeneity hypothesis (accept at the 1.7% level). Again, the tariff coefficient becomes more negative.

It is tempting to conclude that the more negative IV tariff coefficient implies that policy-makers use high wage premia levels as a signal to protect

Table 4
Wage premia regressions with trade and protection instrumented

Dependent variable	Union wage premia (W_i^*)				Non-union wage premia (W_j^*)	
	OLS (1)	IV (2)	OLS (3)	IV (4)	OLS (5)	IV (6)
P: International						
Tariffs	−1.342	−2.476	−1.286	−2.253	−0.158	0.302
	(0.320)	(0.630)	(0.311)	(0.563)	(0.221)	(1.443)
NTBs	0.044	0.104			0.013	0.043
	(0.055)	(0.124)			(0.039)	(0.145)
Imports	−0.348	−0.381	−0.333	−0.344	−0.130	−0.483
	(0.138)	(0.292)	(0.136)	(0.276)	(0.091)	(0.952)
Exports	−0.049	−0.089			0.110	−0.411
	(0.176)	(0.230)			(0.124)	(0.924)
X: Domestic						
Four-firm concentration	0.131	0.135	0.123	0.120	0.125	0.280
	(0.105)	(0.143)	(0.099)	(0.127)	(0.066)	(0.360)
Profit rate	0.144	0.233	0.169	0.276	−0.096	−0.322
	(0.156)	(0.211)	(0.151)	(0.196)	(0.101)	(0.508)
Capital/labor ratio	0.325	0.302	0.329	0.309	0.462	0.387
	(0.228)	(0.255)	(0.225)	(0.245)	(0.150)	(0.230)
Wage decile	0.264	0.198	0.257	0.190		
	(0.085)	(0.103)	(0.084)	(0.099)		
Schooling					0.055	0.080
					(0.014)	(0.044)
MOFA employment	−1.069	3.719	−1.852	1.673	9.864	13.452
	(7.614)	(8.781)	(7.433)	(8.194)	(5.136)	(9.321)
Intercept	−0.447	−0.325	−0.438	−0.318	−0.739	−1.004
	(0.140)	(0.172)	(0.139)	(0.165)	(0.178)	(0.468)
$\bar{R}^2$	0.532	0.494	0.541	0.517	0.506	0.360
Hausman[a]	$\chi_4^2 = 9.01$	(6.1%)	$\chi_2^2 = 8.16$	(1.7%)	$\chi_4^2 = 1.58$	(81.3%)

Note: These are estimates of Eq. (7). Tariffs, imports, NTBs, and exports are instrumented by the domestic regressors as well as factor shares (direct plus indirect in an input-output sense) for two types of capital (plant and equipment; inventories), five occupations (engineers and scientists, white-collar workers, skilled workers, semi-skilled workers, and unskilled workers), three types of land (cropland, pasture, and forest), and three subsoil resources (coal, petroleum, and minerals). Standard errors are in parentheses.

[a] In parentheses is the probability that the trade and protection regressors are exogenous.

an industry. This is mistaken. To illustrate, consider a regression motivated by the government reaction function in Subsection 2.3:

$$\tau_j = W_j^* \gamma_w + m_j \gamma_m + x_j \gamma_x + X_j \gamma_X + u_j \,, \tag{8}$$

where j indexes industries, τ_j is the tariff, W_j^* is the wage premium, m_j is imports, x_j is exports, and X_j are the domestic regressors listed in Table 4. Consider IV estimates of Eq. (8) using the same factor share and domestic instruments as above for W_j^*, m_j, and x_j. As expected, the IV estimate of γ_w is negative (-0.210 with a standard error of 0.059) so that high levels of wage premia lead to low levels of protection.[5]

The non-union sample serves as a control group: to the extent that this group is more heavily represented in competitive product and labor markets we should not expect any of the regressors to be endogenous. As shown in columns (5) and (6), the Hausman test does indeed strongly reject endogeneity. Note also that the tariff coefficient becomes positive, but unlike other results this is not robust to the choice of ω: it only occurs when ω is treated as schooling.

Clearly, we have not exhausted the set of possible instruments. However, our estimates demonstrate that even if a set of instruments were found that led to a positive coefficient for tariffs, the result would depend critically on the choice of instruments and so be of little value. Thus, if the negative coefficient reflects the political–economic impact of wages on tariffs (a very plausible a priori position), then our detailed data on labor markets, trade, and protection are not informative about the effect of tariffs on union wages. On the other hand, our result is consistent with our propositions: although it is an odd finding, it is an important one for designing future research on wages and protection.

4.4. Fixed effects

A possible shortcoming of our cross-section data is that it cannot be used to control for industry fixed effects that may be correlated with our regressors. The existence of such correlations would lead to biased estimates. Lack of NTB data and sufficiently detailed tariff data for years other than 1983 prevent us from constructing a panel. However, two studies that used panel data to estimate the effect of trade flows on wages provide reason for expecting panel data and cross-section data to yield similar estimates. Compare Freeman and Katz's (1991) panel data estimates with

[5] On an econometric note, let β^{IV} and β^{OLS} be the tariff coefficients in Table 4 so that $0 > \beta^{OLS} > \beta^{IV}$. It is straightforward to show that $0 > \text{plim}(\beta^{OLS}) > \text{plim}(\beta^{IV})$ is consistent with $\gamma_w < 0$. That is, the Table 4 results are consistent with the protection of low wage premia industries.

our cross-section specification in column (1) of Table 2. For exports, both studies obtained negative export coefficients: our fragile estimate of −0.05 compares with their estimate of −0.18. For imports, our estimate of −0.35 is almost identical to theirs of −0.37. Thus, similar results obtain even though we have not controlled for fixed effects. The most direct evidence on estimation bias is provided by Macpherson and Stewart (1990) who explicitly examined this issue. They estimated a cross-section regression of wage levels on import levels and a panel regression of wage changes on import changes. While the models are not entirely comparable, the respective import-related coefficient estimates were − 0.0017 and − 0.0015 from which they concluded that "the [panel] results provide general support for the findings from the [cross-section] levels specification discussed earlier" (p. 443). Thus, biases introduced by omitting fixed effects appear to be small.

5. Conclusions

This paper built on the McDonald and Solow (1981) framework by merging Brander and Spencer (1988) preferences with Mezzetti and Dinopoulos (1991) staging and by endogenizing trade policy. The effects of trade, trade policy, and domestic factors on union wages were analyzed and used to frame an empirical study of the role of trade and trade policy in union wage determination. The unique data set combined detailed labor market data with detailed data on trade and protection. Several conclusions emerged. First, trade and protection appear as key variables in the union wage determination process—as important as the domestic regressors considered previously by many researchers. Second, union wage premia are sensitive to import competition, much more so than to export success. This explains union resistance to the Canada–U.S. and Canada–Mexico–U.S. Free Trade Agreements: unions bear the cost of increased import competition, but they do not reap the export-promotion benefits of trade liberalization. Third, the NTB and export coefficients are not robust. Fourth, tariffs are negatively correlated with wage premia.

Clearly, the negative tariff coefficient does not mean that tariff reductions lead to increased wages in the short run. One explanation of the negative tariff coefficient is that policy-makers protect low-wage industries. However, our IV results either rule this out or suggest that the data are not sufficiently informative to disentangle the impact of protection on wages from the impact of wages on protection. Of considerable interest, the endogeneity of protection suggested by the negative tariff coefficient and the endogeneity of trade flows suggested by the strategic trade literature received only weak support. Another explanation of the negative tariff coefficient is the presence of industry fixed effects correlated with both wage premia and tariffs. Evidence from other studies does not support this view, though research using a long time series or panel to control for fixed effects will be necessary to address this hypothesis seriously.

The negative tariff effect is consistent with the theory, a point which suggests a final explanation, namely unions. The ambiguous theoretical effect of trade policy on wages (Propositions 1 and 2) stems from the fact that union workers can negotiate low wages in return for implicit or explicit employment guarantees. This employment–wage trade-off is not available to non-union workers so that a statistically significant negative tariff should not appear for non-union workers. Indeed, tariffs as well as NTBs, imports, and exports are jointly insignificant in the otherwise excellent-fitting non-union wage equation. Thus, there is a fundamental difference between union and non-union wage responses to trade and trade policy.

Data appendix

We merged months from the 1984 CPS that were available from the ICPSR: January, March, May, June, July, October, and November. Only individuals meeting the following criteria were kept in our sample. (1) The individual is in an outgoing rotation group. (2) The individual's hourly wage is between \$1 and \$250, where hourly wage is defined as usual weekly earnings divided by usual weekly hours worked. (3) The individual has positive usual weekly hours. (4) The individual is between 16 and 75 years of age. (5) The individual works in a manufacturing industry that is not described as 'miscellaneous' (CIC industry codes 122, 301, 332, 350, 383, and 392), for which reliable trade data are available (CIC codes 140 and 142 were merged and 362 and 372 deleted) and for which aggregation is sensible (CIC code 211 was deleted). This left us with 5,474 individuals in 68 industries.

Occupations were defined using the CPS occupation classification: Engineers and scientists (43–83); White collar (3–27, 84–389); Skilled (494, 497, 503–699); Semi-skilled (703–859); Unskilled (403–469, 499, 863–889). Workers in farming occupations (472–494) were omitted. 'Lives in South' is defined as living in one of the 17 states in the CPS South Region. Labor force experience is age less highest grade attended less 5.

Data on MOFAs are from *U.S. Direct Investment Abroad, Revised 1983 Estimates*, U.S. Department of Commerce (October 1986), Tables 47(1), 49(1), and 51(5). The data are more highly aggregated across industries than the CPS data and, where necessary, were pro-rated among the CPS industries using U.S. industry output weights. Half of MOFA imports are Auto Pact trade which, given the close link between the Canadian and U.S. auto worker unions, does not confer bargaining strength to the auto makers.

These were excluded from MOFA imports and employment. Including them makes little difference.

Mathematical appendix

Proof of Proposition 1. Define $U_0 = U(\omega)$. Totally differentiating Eqs. (3) and (1), using Eq. (2) to simplify, and stacking yields

$$\begin{bmatrix} U_{ww}(w - R_n) & -U_w R_{nn} \\ a & U - U_0 \end{bmatrix} \begin{bmatrix} \mathrm{d}w \\ \mathrm{d}n \end{bmatrix}$$
$$= \begin{bmatrix} 0 & -1 & 0 & U_w R_{n\rho} \\ n(U - U_0)/\alpha & (1-\alpha)n & -\alpha U_w & \alpha U_w R_\rho \end{bmatrix} \begin{bmatrix} \mathrm{d}\alpha \\ \mathrm{d}U_0 \\ \mathrm{d}\pi \\ \mathrm{d}\rho \end{bmatrix},$$

where $a = nU_w - \alpha U_{ww}(R - wn - \pi)$. Solving for $(\mathrm{d}w, \mathrm{d}n)$ yields

$$\begin{bmatrix} \mathrm{d}w \\ \mathrm{d}n \end{bmatrix} = \frac{1}{C} \begin{bmatrix} U - U_0 & U_w R_{nn} \\ -a & U_{ww}(W - R_n) \end{bmatrix}$$
$$\times \begin{bmatrix} 0 & -1 & 0 & U_w R_{n\rho} \\ n(U - U_0/\alpha) & (1-\alpha)n & -\alpha U_w & \alpha U_w R_\rho \end{bmatrix} \begin{bmatrix} \mathrm{d}\alpha \\ \mathrm{d}U_0 \\ \mathrm{d}\pi \\ \mathrm{d}\rho \end{bmatrix}, \tag{9}$$

where $C = U_{ww}(w - R_n)(U - U_0) + aU_w R_{nn}$ is a determinant. By assumption, $U_w > 0$, $U_{ww} < 0$, $R_{nn} < 0$, $R_\rho > 0$, $U - U_0 > 0$, and $R - wn - \pi > 0$. By Eq. (3), $w - R_n > 0$. Hence $a > 0$ and $C < 0$. The conclusions follow from Eq. (9).

Proof of Proposition 2. Define $U_0 = U(\omega)$. Totally differentiating Eqs. (3), (1), and (4), using Eq. (2) to simplify, and stacking yields

$$\begin{bmatrix} -U_{ww}(w - R_x) & U_w R_{xx} & U_w R_{xy} \\ -b & -(U - U_0) & \alpha U_w R_y \\ 0 & R^*_{yx} & R^*_{yy} \end{bmatrix} \begin{bmatrix} \mathrm{d}w \\ \mathrm{d}x \\ \mathrm{d}y \end{bmatrix}$$
$$\times \begin{bmatrix} 0 & 1 & 0 & 0 \\ -x(U - U_0)/\alpha & -(1-\alpha)x & \alpha U_w & 0 \\ 0 & 0 & 0 & 1 \end{bmatrix} \begin{bmatrix} \mathrm{d}\alpha \\ \mathrm{d}U_0 \\ \mathrm{d}\pi \\ \mathrm{d}\tau \end{bmatrix}, \tag{10}$$

where $b = xU_w - \alpha U_{ww}(R - wx - \pi)$. Solving for $(\mathrm{d}w, \mathrm{d}x, \mathrm{d}y)$ yields

$$\begin{bmatrix} \mathrm{d}w \\ \mathrm{d}x \\ \mathrm{d}y \end{bmatrix} = \frac{U_w}{-D} \begin{bmatrix} (w - R_x)R^*_{yy} + \alpha R_y R^*_{yx} & R_{xx}R^*_{yy} - R_{xy}R^*_{yx} & -\alpha U_w R_y R_{xy} - (U - U_0)R_{xy} \\ -bR^*_{yx}/U_w & U_{ww}(w - R_x)R^*_{yx}/U_w & bR_{xy} - \alpha U_{ww}(w - R_x)R_y \\ bR^*_{yy}/U_w & -U_{ww}(w - R_x)R^*_{yy}/U_w & -U_{ww}(w - R_x)^2 - bR_{xx} \end{bmatrix}$$

$$\times \begin{bmatrix} 0 & 1 & 0 & 0 \\ -x(U - U_0)/\alpha & -(1 - \alpha)x & \alpha U_w & 0 \\ 0 & 0 & 0 & 1 \end{bmatrix} \begin{bmatrix} \mathrm{d}\alpha \\ \mathrm{d}U_0 \\ \mathrm{d}\pi \\ \mathrm{d}\tau \end{bmatrix}, \tag{11}$$

where

$$D = R^*_{yy}U_w(U_{ww}(w - R_x)^2 + bR_{xx}) + R^*_{yx}U_w(\alpha U_{ww}(w - R_x)R_y - bR_{xy}).$$

Sign D using stability. Eqs. (3) and (1) implicitly define the domestic reaction functions $w_{t-1} = S^w(y_t)$ and $x_{t-1} = S^x(y_t)$ and Eq. (4) implicitly defines the foreign reaction function $y_{t-1} = S^y(x_t)$. Then $\partial S^y/\partial x = -R^*_{yx}/R^*_{yy}$, $\partial S^x/\partial y = (\alpha U_{ww}(w - R_x)R_y - bR_{xy})/(U_{ww}(w - R_x)^2 + bR_{xx})$, and $D = U_w(U_{ww}(w - R_x)^2 + bR_{xx})R^*_{yy}(1 - (\partial S^y/\partial x)(\partial S^x/\partial y))$. If $(\partial S^y/\partial x)(\partial S^x/\partial y) < 0$, then $D > 0$. Otherwise, the sufficient condition for stability (Stokey and Lucas, 1989, Theorem 6.5) implies $(1 - (\partial S^y/\partial x)(\partial S^x/\partial y)) > 0$ so that $D > 0$. By assumption $U_w > 0$, $U_{ww} < 0$, $U - U_0 > 0$, $R - wn - \pi > 0$, $R_x > 0$, $R_y < 0$, $R_{xx} < 0$, $R_{xy} < 0$, $R^*_{yy} < 0$, and $R_{xx}R^*_{yy} - R_{xy}R^*_{yx} > 0$. Hence $b > 0$. By Eq. (2), $w - R_x > 0$. The conclusion follows from Eq. (11).

Proof of Proposition 3. Define $U_0 = U(\omega)$, $S^\tau_w = \partial S^\tau/\partial w$, and $S^\tau_x = \partial S^\tau/\partial x$. Alter Eq. (10) by adding a row corresponding to the total differential of the government reaction function, $S^\tau_w \mathrm{d}w + S^\tau_x \mathrm{d}x - \mathrm{d}\tau = 0$, and solve for $(\mathrm{d}w, \mathrm{d}x, \mathrm{d}y, \mathrm{d}\tau)$ in terms of $(\mathrm{d}\alpha, \mathrm{d}U_0, \mathrm{d}\pi)$. Let E be the determinant of the 4×4 left-hand-side matrix. The key results are $\mathrm{d}w/\mathrm{d}\alpha = -cxU_w(U - U_0)/(\alpha E)$, $\mathrm{d}w/\mathrm{d}\pi = c\alpha U^2_w/E$, and $\mathrm{d}w/\mathrm{d}U_0 = -cxU_w(1 - \alpha)/E + (U - U_0)R^*_{yy}/E + \alpha U_w R_y(R^*_{yx} - S^\tau_x)/E$, where $c = R_{xx}R^*_{yy} - R_{xy}R^*_{yx} + S^\tau_x R_{xy}$.

By assumption, $R_{xx}R^*_{yy} - R_{xy}R^*_{yx} > 0$, $S^\tau_x < 0$, and $R_{xy} < 0$. Hence $c > 0$. It remains to sign E. Eqs. (3), (1), and (4) define the private sector reactions of w, x, and y to τ. Write these as $S^w(\tau)$, $S^x(\tau)$, and $S^y(\tau)$. Then $E = -D(1 - S^\tau_w S^w_\tau - S^\tau_x S^x_\tau)$, where $D > 0$ is defined in the proof of Proposition 2. If $S^\tau_w S^w_\tau + S^\tau_x S^x_\tau < 0$, then $E < 0$. Otherwise, the sufficient condition for stability of the private sector–public sector game (Stokey and Lucas, 1989, Theorem 6.5) implies $(1 - S^\tau_w S^w_\tau - S^\tau_x S^x_\tau) > 0$ so that $E < 0$.

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