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Multinational Corporations, Wages and Employment: Do Adjustment Costs Matter?

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Abstract

This paper investigates the extent to which expansion of international production by US multinationals reduces labor demand at home and at other foreign locations. The few existing empirical studies on this subject maintain, along with …rm's cost minimizing behavior, perfectly variable inputs. In our view, this constitutes a strong restriction, which may seriously bias the short run cross-price elasticity estimates. We o¤ers simple examples in which this is actually the case. We suggest an alternative approach to the problem, which explicitly considers for the presence of adjustment costs. A dynamic model of the …rm is applied to estimate short-run and long-run cross-price elasticities between home and foreign labor. We …nd evidence of signi…cant adjustment costs for employment in Latin American and Canadian a¢liates. We also …nd that, in some instances, due to the presence of slow input adjustments, the complementarity/substitution relationship between employment in di¤erent international locations is reversed from the short to the long-run.

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Introduction

In the ongoing debate on the effects of globalization on wages and employment, one of the most crucial issues is the role played by multinational companies. It is claimed that, as long as multinationals establish and/or expand overseas production, they tend to substitute workers at home with workers in foreign affiliates' countries. Then one of the most studied aspect of home country effects of globalization has been whether employment in foreign affiliates is a 'substitute' or a 'complement' to home country employment of the parent firms.

This paper tackles substitution-versus-complementarity issue by applying an empirical model of multinational behavior that accommodates input adjustment costs. In fact, the presence of costs in adjusting quasi-fixed inputs influences factor demand decisions and the related price elasticities, and consequently explanations of firm behavior using static model may be misleading. Our results confirm that this is actually the case. In some instances, due to the presence of slow input adjustments, the complementarity/substitution relationship between employment in different international locations is reversed from the short to the long-run.

In the recent literature, a number of authors have investigated the interplay of multinationals' activities and labor markets by using static approach. Slaughter (1995), using industry-distributed data, finds weak evidence of US multinationals' substituting home with foreign labor. Brainard and Riker (1997a), (1997b) confirm Slaughter's finding for a panel data of US foreign affiliates and parents. Slaughter (1995), using Bureau of Economic Analysis (BEA) data on US manufacturing multinationals in the 1980s, estimates multinationals' cost-minimization functions which treat all factors of production, both domestic and foreign, as jointly chosen. From these functions, price elasticities of demand between US and foreign labor are estimated. He chooses a translog specification of the cost function and finds that domestic industry employment and overseas affiliate employment are complementary, but only weakly so. Brainard and Riker (1997a) use a firm-level panel of US multinationals and their foreign affiliates between 1983 and 1992. They analyze labor demand within firms across plant locations, by fitting a firm-level global cost function specified in terms of relative wages (which are taken as exogenous). They aggregate labor across subsets of countries and treat each subset as separate factors. Brainard and Riker estimate substitution among locations conditional on the capital stock. The authors find that substitution between labor employment by parents in the United States and affiliates abroad is low. In contrast, in a related paper, they find that there is strong substitution between workers at affiliates in alternative low wage locations, where the activities most sensitive to labor costs are performed (Brainard and Riker, 1997b). In a recent paper on Swedish multinationals, Braconier and Ekholm (1999) find similar results.

Another strand in the literature has investigated the role of multinationals' activities by focussing on whether production abroad tends to raise or lower the labor intensity of home production. In Blomstrom, Fors and Lipsey (1997), US multinationals seem to show lower labor intensity of home production than Swedish parents in the presence of higher foreign production. The authors' argue that this result reflects the fact that US firms, much more than Swedish firms, invest abroad to take advantage of factor price differences in a vertical division of activities among countries. footnote Similar conclusions are presented by Bassino (1998), studying the effects on domestic labor market of increasing foreign direct investments towards low-wage Eastern Asian countries by Japanese multinationals during the recession of the 1990s. footnote

A relevant question is whether multinationals' behavior is more accurately modeled using a dynamic specification or adequately represented by static approaches. All the studies dealing with the substitution-versus-complementarity issue maintain, along with firm's cost minimizing behavior, perfectly variable inputs and perfectly competitive labor markets. In our view, this static framework is not adequate to set up an empirical model of multinationals' behavior. We show by a simple example that in a world in which employment adjustment costs and labor market imperfections are important, a standard cost function approach can lead to misleading evidence about the substitution-versus-complementarity issue.

In this paper, we suggest and discuss an alternative line of attack to the problem, which explicitly allows for the presence of adjustment costs. We measure short-run and long-run cross-price elasticities within a more general set-up, in which the firm chooses variable inputs and quasi-fixed inputs profiles in order to minimize its discounted flow of total cost over a given time horizon. Also, we test for the presence of significant adjustment costs. This model was first formalized by Treadway (1971) using a primal approach and then by Epstein (1981), using a dual approach. The first empirical application of the dual approach is due to Epstein and Denny (1983).

In this study on US multinationals, we treat labor in different geographical locations as quasi-fixed inputs and we test for the presence of adjustment costs. Although there is some disagreement as to the structure and the size of the cost of adjusting factor demand, the literature agrees that adjustment costs arising within a firm lead to optimal multi-period modifications of the levels of quasi-fixed inputs in response to single-period price changes; and that a firm suffers short-run costs of changing stocks of quasi-fixed factors. footnote The sources of adjustment costs associated with changing employment may be related to: searching activities (i.e. screening and processing new employees); training (including disruptions to production as previously trained workers' time is devoted to on-the-job instruction of new workers); severance pay (mandated and otherwise). When multiplant firms are considered, as is the case of multinational corporations, the closing and opening of plants generates net adjustment costs.

The presence of costs in adjusting quasi-fixed inputs influences factor demand decisions and the related price elasticities, and consequently explanations of firm behavior using static model may be misleading. We estimate short-run and long-run cross-price elasticities of labor demand in different locations, and we find that the relationship of complementarity or substitution between the employment in different geographical areas may be reversed due to the presence of significant adjustment costs.

The structure of the paper is as follows. In the next section it is shown that problems of incorrect inference can arise within the standard static cost function approach to multinationals' international production. Section 4, after discussing some stylized facts on US multinationals (Section 3), illustrates an alternative method of empirical analysis - the dynamic duality approach - which may overcome some of the problems associated with the standard approach. The estimation and testing strategy is presented in Section 5. Finally, in Section 6, the results are discussed.

Cost Minimizing Behavior and Price Substitutability: A Simple Example

In what follows our attention is centered upon the cost minimization approach. We see several sources of misspecification arising in such a context. To begin with, capital and employment adjustment costs are likely to be not negligible, both in parent and affiliate firms. Slaughter (1995) and Brainard and Riker (1997a) actually allow for both parent and affiliate capital as quasi-fixed inputs, by conditioning the variable cost function on the capital stock. We claim that this is not enough if labor is quasi-fixed because of adjustment costs in terms of foregone output. In this case the employment variation with respect to the past year, and not just the current employment, is likely to have an impact on the variable costs. Neglecting this possibility may therefore seriously bias the estimates. Finally, non-competitive labor markets may also play a role as an important source of misspecification.

As a prelude to the formal analysis of cost minimizing behavior in the presence of costly adjusted inputs, we present a simple example of how, in such instance, a static optimization approach can lead to incorrect inference about the nature of the input substitutability. This example constitutes a simplified version of the general model upon which our empirical application is based. As such, it provides a convenient introduction to some of the most crucial aspects of the empirical analysis.

Consider a multinational company that produces a single output y from costly adjusted employment at region A and perfectly variable employment at region B. The multinational company minimizes its discounted flow of costs over an infinite horizon. Assume for simplicity that $y = 1$

always. Also, assume that the intertemporal cost minimization problem yields the following employment demand function in region *A*

$$l = L + K\tilde{w} + Jl_{-1} \qquad \#$$

where $\tilde{w}$ denotes the wage paid to workers at *A* normalized by the wage paid to workers at *B*, and l_{-1} denotes the employment at *A* lagged one time. Optimizing behavior implies $K < 0$. In addition, $0 < J < 1$ and $L + K\tilde{w} \geq 0$, which together imply that the employment path described by (ref: affpath) converges to a unique stable steady state

$$\bar{l} = (L + K\tilde{w})/(1 - J). \qquad \#$$

Now, it is clear from (ref: steady) that unless l is always at its steady state level, there is no hope to infer the long-run wage impact $K/(1 - J)$ by regressing l on $\tilde{w}$ alone. Nor, is it generally possible to infer the short-run behavior K from the static regression. Let us see why this is so. If the current wage and the past employment are correlated, as is likely in practice, the coefficient K would not be identified in a regression that does not include l_{-1} on the right. Contract bargaining with a monopoly union is one simple instance in which $\tilde{w}$ depends on l_{-1}. Assume that while the multinational company takes $\tilde{w}$ as given and determines l according to (ref: affpath), $\tilde{w}$ is chosen by a union with the following objective function

$$u(\tilde{w}, l) \qquad \#$$

The function u is increasing in both arguments, strictly concave in l and such that $\frac{\partial^2 u(\tilde{w}, l)}{\partial \tilde{w} \partial l} \geq 0$. The union chooses $\tilde{w}$ to maximize (ref: union) subject to (ref: affpath). We assume the existence of an internal solution $\tilde{w}(l_{-1})$, which satisfies the first order condition

$$u_1 - \frac{K}{\tilde{w}} u_2 = 0 \qquad \#$$

By applying the Implicit Function Theorem to the first order condition of such optimization problem and using the second order condition yield the following comparative static result

$$\frac{\partial \tilde{w}(l_{-1})}{\partial l_{-1}} \geq 0$$

Therefore, a rise in the past employment causes both a rise in the current compensation, through the union bargaining mechanism and a rise in the current employment, through the adjustment costs effect. This is the crucial point of our example. If the role of the lagged employment in the employment demand function is neglected, comovements of $\tilde{w}$ and l, induced by lagged employment variation across multinational firms, are mistaken as labor at *A* and labor at *B* being price complements footnote .

Of course, this is not the only case in which we might find an association between current wage and past employment. In fact, in the presence of both nominal and real rigidities, wages respond to unexpected shocks slower than employment.

Finally, it is worth noting that the above considerations, besides emphasizing the necessity of including input adjustment costs into the model specification, also suggest to treat the wage as an endogenous variable. In fact, whatever type of labor market imperfection occurs in practice, it is however very likely that $\tilde{w}$ depends on any random shock affecting the employment demand, casting some doubt on the regression analyses that treat $\tilde{w}$ as exogenous.

With this in mind, we believe that more appropriate analytical tools than static cost minimization should be employed. In particular, techniques allowing for the presence of adjustment costs seems particularly useful in the analysis of multinationals' international production. Also, employing instrumental variable based estimation methods should help in attenuating the bias induced by labor market imperfections. These are exactly the directions along which we shall move for the empirical application.

Data and Stylized Facts on US Multinationals

This study uses industry-level data on US multinational companies. The data are collected and administered on a mandatory basis by the Bureau of Economic Analysis (BEA). The BEA Annual and Benchmark Surveys of US Direct Investment Abroad report foreign affiliates and US parents financial and operating data on an annual basis. footnote

The data set includes all parents and foreign affiliates classified in the manufacturing sector. The activities of manufacturing multinationals are aggregated into 32 different industries. We use data from 1982 to 1994 and construct a panel including parents and affiliates total employment, employee compensation, sales and total assets. footnote Moreover, foreign affiliates are grouped into four geographical areas reflecting, in a very broad sense, proximity and development differences: Canada, Europe, Latin America and Rest of the World.

Before presenting the econometric estimates, we illustrate some stylized facts on US multinationals. First we check whether there is some evidence on substitution between labor at home and labor in foreign affiliates, particularly in low wage countries.

We find that, over the period 1982-1994, employment is declining both in parents and in foreign affiliates, but in percentage terms this decline was larger for parents (-13.5%) than for affiliates (-7.1%). As a result, relative employment in multinationals shifted slightly towards affiliates (Figure 1).

Looking at how the allocation of affiliates' employment has evolved across locations over time, Figure 2 shows that the share of employment at affiliates in industrialized countries has marginally declined, while the share at affiliates in Latin America has slightly expanded until 1993, with a decline in 1994. On the contrary, the share of the Rest of the World shows a slow growth until 1993, and then an increase, offsetting the decline in Latin America.

Figure 3 shows corresponding changes in relative wages, defined as employee compensation of affiliates in different locations over employee compensation of the parent. Figures 2 and 3 seem to suggest that there is no simple correlation between changes in employment share and changes in relative wages. In fact, increasing relative wages, as in the case of locations in Europe or in the Rest of the World, are not associated to declining employment share and, at the same time, low relative wages as in Latin America seem not to cause employment expansion.

Finally, we look at the persistence of employment levels over time, to find some evidence for the presence of adjustment costs. We calculate the autocorrelation coefficients of employment's annual changes of the parents and of foreign affiliates in different geographical areas. We find that foreign affiliates in Latin America and in Canada have positive degrees of persistence in employment levels (respectively 0.17 and 0.11). Differently, autocorrelation coefficients for Europe, the Rest of the World and the parents are negligible, suggesting absence of adjustment costs (Table 1).

The structural econometric model is going to shed light on the significance of the foregoing stylized facts.

The Dynamic Duality Approach to Adjustment Costs

The Theory

The theory of adjustment cost in production postulates that firms may suffer a short run loss in terms of foregone output when they change their stocks of quasi-fixed inputs. To bring adjustment costs into the analysis of multinational behavior, we follow the dynamic duality approach as developed and implemented by Epstein and Denny (1983). The dual approach has the distinctive advantage of simplifying the derivation of both the closed form solutions for the optimal input demands and the complete list of model empirical implications.

The theory assumes that at time t the firm combines variable inputs $v(t)$ and quasi-fixed inputs $l(t)$ to produce output $y(t)$ according to the production function footnote

$$y = F(\dot{l}, l, v) \qquad \#$$

Adjustments costs in terms of foregone output are modelled by assuming that $F_{\dot{l}} > 0$ if $\dot{l} < 0$ and $F_{\dot{l}} < 0$ if $\dot{l} > 0$ footnote . From F, we define the variable cost function C

$$C(\dot{l}, l, p_v, y) = \left[\min_v p_v v \, : \, y = F(\dot{l}, l, v)\right],$$

where p_v is the vector of variable input prices. The firm is price taker with respect to p_v and the quasi fixed input prices p_l, which are assumed always strictly positive. Let $r > 0$ denote the real rate of discount, assumed constant over time. It is also assumed that the firm holds static expectations on p_l, p_v and y. Given this, the firm chooses the *rate of change* path of the quasi fixed inputs, $\dot{l}$, (dots denote derivatives with respect to time) and the variable input level v to minimize its discounted flow of total costs (quasi-fixed cost added to variable cost) in continuous time and over an infinite horizon. This defines the following intertemporal optimization problem with the associated optimized value function

$$J(l, p_l, p_v, y) = \min_{\dot{l}, v} \int_0^{+\infty} e^{-rt}[p_l l + p_v v] dt \qquad \#$$

subject to the technological constraint

$$y = F(\dot{l}, l, v)$$

and $l(0) = l_0 > 0$.

Using the variable cost function, the problem (ref: minval) can be reformulated in a more compact form

$$J(l, p_l, p_v, y) = \min_{\dot{l}, v} \int_0^{+\infty} e^{-rt}\left[p_l l + C(\dot{l}, l, p_v, y)\right] dt \qquad \#$$

subject to $l(0) = l_0 > 0$, [see Epstein and Denny (1983)].

The variable cost function C offers a representation of the technology that is equivalent to F. Thus, we assume that F is such that C satisfies the following properties a) $C \geq 0$; b) $C_l < 0$; c) $C_{\dot{l}} > 0$ if $\dot{l} > 0$ and $C_{\dot{l}} < 0$ if $\dot{l} < 0$; d) C is convex in $\dot{l}$; e) the problem (ref: MinVal1) (or (ref: minval)) has a unique globally stable steady state. Notice that property c) is dual to the adjustment cost property of F.

The function J has a useful property, namely it satisfies the Hamilton-Jacobi equation [see Arrow and Kurz (1970)]:

$$rJ(l, p_l, p_v, y) = \min_{\dot{l}}\left[p_l l + C(\dot{l}, l, p_v, y) + J_l(l, p_l, p_v, y)\, \dot{l}\right]$$

where J_l is the vector of shadow prices associated with quasi-fixed stocks. The Envelope theorem and Shephard Lemma can then be applied to the foregoing equation to derive the policy functions for the quasi fixed input rates of change, $\dot{l}^*$, and the variable input demands, v^*, solutions to (ref: minval):

$$\dot{l}^*(l, p_l, p_v, y) = J_{lp_l}^{-1}[rJ_{p_l} - l] \qquad \#$$

$$v^*(l, p_l, p_v, y) = r(J - J_{p_l} p_l) - (J_l - p_l J_{lp_v})\, \dot{l}^*$$

Furthermore, the duality between C and J uncovers the J must satisfy the following list of regularity conditions: a) $J \geq 0$; b) $rJ_l - p_l - J_{ll}\, \dot{l}^* < 0$; c) $J_l < 0$ if $\dot{l}^* > 0$ and $J_l > 0$ if $\dot{l}^* < 0$; d) $rJ_y - J_{yl}\, \dot{l}^* > 0$; e) $rJ - J_l\, \dot{l}^*$ is concave in p_l; f) the stock profile l^* associated with $\dot{l}^*(l, p_l, p_v, y)$ and l_0 has a unique globally stable steady state $\bar{l}$. Such conditions also offer an exhaustive characterization of J [see Epstein and Denny (1983)].

The equations in (ref: ddsystem) constitute the system upon which we base our empirical analysis.

The Empirical Model

We assume that a US multinational can allocate labor among the home parent and its affiliates throughout the world. We focus on four different areas: Canada, Europe, Latin America and the Rest of the world. Therefore, five production inputs are considered: parent employment, l^p, affiliate employment in Europe, l^e, affiliate employment in Canada, l^c, affiliate employment in Latin America, l^l and affiliate employment in the rest of the world, l^{rw}. For each input we have the correspondent input price: p_{l^p}, p_{l^e}, p_{l^c}, p_{l^l} and $p_{l^{rw}}$.

We maintain throughout that l^{rw} is the only variable input, while l^p, l^e, l^c and l^l are quasi fixed. Thus $v \equiv l^{rw}$, $p_v \equiv p_{l^{rw}}$, $l \equiv \{l^p, l^e, l^c, l^l\}$ and $p_l \equiv \{p_{l^p}, p_{l^e}, p_{l^c}, p_{l^l}\}$.

As in many dynamic duality applications, we choose a quadratic specification for J (see Epstein and Denny (1983) among others).

$$J = A_o^i + A_{p_l}^i p_l + A_{p_l y} p_l y + A_{p_v} p_v + A_l l + \quad \#$$

$$+ 1/2\left[y p_l^T A_{p_l p_l} p_l + p_v^T A_{p_v p_v} p_v + p_l^T A_{p_l l}^{-1} l + p_v^T A_{p_v l}^{-1} l + p_l^T A_{p_l p_v} p_v + l^T A_{ll} l\right]$$

$A_o^i, A_{p_l}^i, A_{p_l y}, A_{p_v}, A_l$ are (4×1) column vectors and $A_{p_l p_l}, A_{p_v p_v}, A_{p_l l}^{-1}, A_{p_v l}^{-1}, A_{p_l p_v}, A_{ll}$ are (4×4) matrixes of constant parameters. We accommodate individual heterogeneity by allowing A_o^i and $A_{p_l}^i$ to vary across sectors.

Estimating the whole system (ref: ddsystem) would require the estimation of an extremely large number of parameters, compared to the size of BEA data set. We, therefore, restrict ourselves to the quasi fixed input subsystem. Applying (ref: ddsystem) to (ref: quad), it has

$$\dot{l}^D = A_{p_l l}\{A_{p_l}^i + ry(A_{p_l y} + A_{p_l p_l} p_l)\} + (rI - A_{p_l l})l. \quad \#$$

To make (ref: qfsyst) empirically implementable, we follow a standard practice in the continuous time investment literature, which considers the discrete time approximation of (ref: qfsyst):

$$l = A_{p_l l} A_{p_l}^i + ry A_{p_l l}(A_{p_l y} + A_{p_l p_l} p_l) + (rI + I - A_{p_l l}) l_{-1}, \quad \#$$

where I is the (4×4) identity matrix. The reader can also notice that the model yields sector-specific coefficients $A_{p_l l} A_{p_l}^i$.

For future reference, let

$$A_{p_l y} = \begin{pmatrix} K_p \\ K_e \\ K_c \\ K_l \end{pmatrix},$$

$$A_{p_l p_l} = \begin{pmatrix} K_{pp} & K_{pe} & K_{pc} & K_{pl} \\ K_{pe} & K_{ee} & K_{ec} & K_{el} \\ K_{pc} & K_{ec} & K_{cc} & K_{cl} \\ K_{pl} & K_{el} & K_{cl} & K_{ll} \end{pmatrix}$$

and

$$A_{p_l l} = \begin{pmatrix} J_{pp} & J_{pe} & J_{pc} & J_{pl} \\ J_{ep} & J_{ee} & J_{ec} & J_{el} \\ J_{cp} & J_{ce} & J_{cc} & J_{cl} \\ J_{lp} & J_{le} & J_{lc} & J_{ll} \end{pmatrix}.$$

According to the regularity conditions spelled out for J, $A_{p_l l}$ does not need to be a symmetric matrix.

Estimation and testing strategy

Estimation

We now turn to the econometric methodology, starting with the estimation issues. To capture non systematic optimization errors and unobservable industry and time heterogeneity, we append to the system (ref: qfsyst) the vector of random variables $\upsilon \equiv [\upsilon^{p\prime}, \upsilon^{c\prime}, \upsilon^{e\prime}, \upsilon^{la\prime}]^{\prime}$. We maintain $E(\upsilon^h_{it}) = E(\upsilon^h_{it}\upsilon^{r\prime}_{js}) = 0$ for all $t \neq s = 1,...,T_i$, all $h \neq r = p, c, e, la$ and all $i \neq j = 1,...,N$. Let $\upsilon_{it} \equiv [\upsilon^p_{it}, \upsilon^c_{it}, \upsilon^e_{it}, \upsilon^{la}_{it}]^{\prime}$, we also maintain $E(\upsilon_{it}\upsilon^{\prime}_{it}) = \Sigma_{it}$, for all $i = 1,...,N$ and $t \neq s = 1,...,T_i$. That is, while we impose lack of serial correlation, we allow for simultaneous cross-equation correlation. Quasi fixed input prices are normalized by the price of the variable input p_v.

It is well known from the dynamic panel data literature that the standard within estimator applied to a first order autoregressive model yields inconsistent estimates when only the number of time periods T_i is large [see Nickell (1981), among others], which is not the case for most panels. In our application the number of sector N is fixed footnote , but it is nonetheless large with respect to T_i and, so, the inconsistency problem could well arise. To solve such a problem, econometricians have suggested various instrumental variables approaches [Anderson and Hsiao (1982), Arellano and Bond (1991), Ahn and Schmidt (1995) among others]. Here we follow the Generalized Method of Moment estimator approach suggested by Arellano and Bond, widely used in most recent dynamic panel data applications [Blundell, Bond, Devereux and Schiantarelli (1992); Nickell, Wadhwani and Wall (1992); Bond and Meghir (1994); Galeotti, Guiso, Sack and Schiantarelli (1996) among others], which exploits all available linear orthogonality conditions. The procedure goes as follows.

The individual effects are eliminated by taking the model in first differences:

$$\Delta l = r[A_{p_l l} A_{p_l y} \Delta y + A_{p_l l} A_{p_l p_l} \Delta(y p_l)] + (rI + I - A_{p_l l})\Delta l_{-1} + \Delta\upsilon \qquad \#$$

Maintaining lack of serial correlation among the errors in levels does not rule out, however, the possibility that the errors in first-differences exhibit MA(1) autocorrelation. We, therefore, follow an instrumental variables approach to estimate (ref: fdiff), with the instruments optimally weighted by the expected variance-covariance matrix of the orthogonality conditions, as required by an optimal Generalized Method of Moment estimator [see Hansen (1982); Gallant (1987), p. 442].

The model generates valid instrumental variables under the form of opportunely trimmed lagged endogenous variables. In fact, as υ is a vector of serially uncorrelated random variables, the following orthogonality conditions hold: $E(\Delta\upsilon^j_{it}\ l^j_{it-s}) = 0,\ s = 2,...,t-1,\ t = 3,...,T_i,\ i = 1,...,N$ and $j = p,\ e,\ c$ and l. Only these orthogonality conditions are exploited for estimation.

As a result, $l_{i,-2}$, $l_{i,-3}$, and further lags yield valid instruments. This is obviously true for all of the endogenous variables considered. So, on principle, $(T_M - 1)(T_M - 2)/2$ instruments can be obtained from each endogenous variable, where $T_M = max[T_1,...,T_N]$. For example, if $N = 2,\ T_1 = 5$ and $T_2 = 6$, we have the following matrix of instruments Z generated by $l_{1,-2}$, $l_{1,-3}$, $l_{1,-4}$, $l_{2,-2}$, $l_{2,-3}$, $l_{2,-4}$, and $l_{2,-5}$

$$Z = . \begin{pmatrix} l_{11} & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & l_{11} & l_{12} & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & l_{11} & l_{12} & l_{13} & 0 & 0 & 0 & 0 \\ l_{21} & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & l_{21} & l_{22} & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & l_{21} & l_{22} & l_{23} & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & l_{21} & l_{22} & l_{23} & l_{24} \end{pmatrix}$$

As already mentioned in Section2, another source of departure of this analysis from the received literature on multinationals is that we treat output y and relative wages p_l/p_v as potentially endogenous. In fact, output is a strategic variable for a multinational firm, as such it could as well be chosen along with the employment variations. It is also very likely that wage, employment and output are affected by correled random shocks. In the view of these considerations, we instrument both y and p_l/p_v by using the appropriate lags.

Capital stock, measured as total assets of the firm, is added to each equation in level and also included into the instrument set with the appropriate lags. To capture world economy effects, we add time period dummies to each equation and also employ them as instruments. Then, abiding by a standard practice in the dynamic duality applications, we take the real discount rate to be equal to a given plausible value. The results reported are for $r = 0.07$. Other values have been tried, confirming all of the results obtained for $r = 0.07$.

Our model is a dynamic four-equation system, using a potentially high number of instruments. As such, it is computationally troublesome. In order to keep the computational difficulties to a minimum, we have restricted our instrument set to the third and fourth lags for each endogenous variable. Besides the other advantages, such choice of the instruments also permits to wipe out measurement errors in the differenced system that present a degree of serial correlation up to the second order.

Finally, the model is estimated by applying a Generalized Method of Moment estimator, as coded by the TSP 4.4 procedure GMM, Following Bruno (1999), we adapt the single equation two-step procedure by Arellano and Bond (1991) to a system of equations with cross-equation parameter restrictions. This procedure is briefly summarized in Appendix A.

Hypothesis testing

Turning to the testing aspects, a convenient feature of our model is that it nests the case of perfectly variable inputs. This hypothesis, therefore, can be tested by computing the joint significance of the set of implied parameter restrictions. The reader can easily verify that the system in (ref: qfdiscr) can be re-arranged according to the following more convenient form

$$\Delta l = M\left[l_{-1} - \bar{l}(p_l, y) \right] \qquad \#$$

where $M = rI - A_{p_l l}$ is the adjustment matrix

$$M = \begin{pmatrix} r - J_{pp} & -J_{pe} & -J_{pc} & -J_{pl} \\ -J_{ep} & r - J_{ee} & -J_{ec} & -J_{el} \\ -J_{cp} & -J_{ce} & r - J_{cc} & -J_{cl} \\ -J_{lp} & -J_{le} & -J_{lc} & r - J_{ll} \end{pmatrix}, \qquad \#$$

and $\bar{l}(p_l, y) = -(rI - A_{p_l l})^{-1}\{A_{p_l l}A^{i}_{p_l} + ryA_{p_l l}(A_{p_l y} + A_{p_l p_l} p_l)\}$ is the vector of steady state stock demands. The system (ref: flacc) expresses the short run input adjustment as a proportion of the desired change.

The parameter restrictions implied by the assumption that, say, l^p is perfectly variable are readily obtained:

$$r - J_{pp} = -1, \; J_{ep} = 0, \; J_{cp} = 0, \; J_{lp} = 0. \qquad \#$$

The interpretation of Conditions (ref: variab) is straightforward. A given input is perfectly variable if a) its short run change is always a 100% adjustment to the desired change, when the other inputs are at their steady state levels and b) the size of its desired change does not interact with the short run adjustments of the other inputs. Conditions (ref: variab) can be tested by using standard Wald tests on the unrestricted matrix M.

As to the regularity conditions for J, the estimation of system (ref: flacc) (or ref: fdiff) allows us to check conditions e) and f). Condition e) boils down to $A_{p_l p_l}$ being a symmetric and negative-definite matrix, while condition f) implies that $M + I$ is a stable matrix.

Results

We describe our empirical results in two sections. First, we present the estimation and testing results from three model specifications. In particular, we check for the significant presence of input adjustment costs and we also asses the fulfillment of the testable regularity conditions. Finally, to asses the issue of input substitutability, we report short run and long run elasticity estimates.

Estimation and Testing Results

The panel based on BEA data set covers the period 1982-1994 for 32 sectors. Due to confidentiality concerns some observations are missing, which attributes to the data set an unbalanced structure. The panel is unbalanced both because the number of observations is different across industries, and because the observations for the industries do not overlap over time. For 3 sectors there are no years without missing data, as a result these sectors are lost in the data set we use. Besides, after constructing lags and first differencing, the first two observations of any uninterrupted interval spanning less than three years are lost for each sector. This makes other 4 sectors disappear, eventually leaving us with 25 sectors. However, there are still enough moment restrictions available to estimate the whole set of parameters.

For any specification tried we report and comment the two-step estimates. The findings are however qualitatively the same in the two cases.

We start by presenting the results from the complete parametric specification (Model 1). Parameter estimates for Model 1 are shown in Table 2 footnote . We have 186 degrees of freedom corresponding to the number of over-identifying restrictions exploited minus the number of estimated parameters. Both the one-step and two-step Sargan-Hansen test of the over-identifying restrictions do not reject the validity of the instruments used at any conventional significance level.

Before proceeding further, a crucial issue is whether the selection rule determining the unbalanced structure of the panel depends or not on the variables to explain. If it does, then ignoring the selection rule would lead to biased estimates. We tackle this point by adopting a simple variable addition test (Verbeek and Nijman,1997). Roughly, this test is based on the idea that if the selection rule is not ignorable then the pattern of missing observations should significantly affect the relationship between endogenous and exogenous variables. Thus, Verbeek and Nijman suggest to include three variables supposedly comprising the effect of the missing data: T_i, the number of years sector i does not present missing data, c_i a dummy variable that is equal to one if sector i never presents missing data and, finally, $s_{i,t?1}$ the selection indicator lagged one time, indicating whether sector i is observed in the previous year. If these variables are significant, then there is evidence of non-ignorable selection rules and a correction procedure for sample selection bias should be adopted.

A problem with this approach is that the first two variables proposed are in fact time-invariant and so the corresponding coefficients are not identified in a fixed effects framework. Moreover, as the dynamic panel approach imposes to discard any uninterrupted interval spanning less than three years, all the 25 sectors are in fact observed in the previous two periods, which makes both $s_{i,t?1}$ and $s_{i,t?2}$ time-invariant. We therefore apply a modification of the Verbeek and Nijman approach to Model 1, which relies on $s_{i,t?3}$ and the interaction between $s_{i,t?3}$ and T_i. We find that, both in the one-step and in the two step estimation, the Wald test of joint significance of the two variables in the four equations does not reject the hypothesis of ignorability at any conventional level of significance: the one-step procedure yields $e^2Ý8Þ = 6.829$, ($P ? value = 0.555$), and the two-step procedure yields $e^2Ý8Þ = 11.687$ ($P ? value = 0.166$). Therefore, even if this finding is by all means not conclusive about the absence of sample selection bias, we proceed further without considering sample selection in our estimation procedures.

None of the testable regularity conditions can be rejected. As to condition f), $rI + I ? A_{p_l l}$ is a stable matrix, as all its eigenvalues turn out to be within the unit circle, which implies that the employment profiles converge to a unique steady state. The other regularity condition is e), which imposes that $A_{p_l p_l}$ is a symmetric and negative definite matrix. We maintain symmetry and test for the negative definiteness. Results are less clear-cut in this case as some principal minors have the wrong sign. However, none of the wrongly signed principal minors turns out to be significantly different from zero.

It is worth analyzing both the size and the statistical significance of the adjustment speeds in all areas. To begin with, let us focus on the own adjustment speeds, which are given by the coefficients on the main diagonal of the adjustment matrix M in (ref: M). They measure the actual input variation within one year as a proportion of the desired change in the input stock. The results are quite consistent with the stylized facts shown in Section 3. First, adjustment costs are important in the case of affiliate employment in Latin America and Canada. In fact, if the other inputs are at their steady state levels, then only 69% of the adjustment to any desired change in the stock of Latin America affiliate employment occurs within one year. A similar result obtains for affiliate employment in Canada, where only 78% of the adjustment occur within one year. Secondly, adjustments are much faster for both parent employment and Europe affiliate employment: very close to 100% within one year.

However, to properly assess whether an input is perfectly variable we need to test the joint parameter restrictions in (ref: variab), involving not only the own adjustment speed, but also the interaction terms in the other input demands. In this respect, tests of perfect variability are quite clear-cut. On the one hand, we can reject the assumption of all inputs being perfectly variable at any conventional significance level, as the Wald test on the joint parameter restrictions (ref: variab) for all inputs is $\chi^2(16) = 31.72$ (p-value = 0.010). On the other hand, we fail to reject the same hypothesis when restricted to parent employment and Europe affiliate employment, since the Wald test on the joint restrictions (ref: variab) for this pair of inputs takes on the value $\chi^2(8) = 7.59$, (p-value = 0.474).

In Model 1 a large number of parameters are not individually significant, as results from Table 2. In addition, a Wald test of joint significance on 40 such parameters yields $\chi^2(40) = 25.97$ with a probability value of 0.957, which does not reject the joint restrictions. Therefore, to obtain a more parsimonious specification, we have set these parameters to zero and estimated the system (ref: fdiff) with such restrictions imposed (Model 2).

The parameter estimates for Model 2 are displayed in Table 3. We now have 226 degrees of freedom and the value of the Sargan-Hansen test of the overidentifying restrictions is 164.10, with a probability value of 0.999 to strongly support the validity of our instruments.

As in Model 1, none of the testable regularity conditions is significantly refuted. As to adjustment speeds, we do not register dramatic changes with respect to the evidence from Model 1. Affiliate employment in Latin America still has the slowest adjustment, 76% of the desired change within one year, even if quite close to the speed of affiliate Canada employment: 77%. Again, adjustment speeds for both parent employment and affiliate employment in Europe are not significantly different from -1, which is evidence of a 100% adjustment.

Our third specification (Model 3), besides the restriction of Model 2, also maintains that parent employment and affiliate employment in Europe are perfectly variable inputs. Restrictions (ref: variab) are therefore imposed for both such inputs. The Wald test carried out within Model 1 for the complete set of restrictions imposed in Model 3 does not reject those restrictions ($\chi^2(44) = 31.25$ with a probability value of 0.925). Table 4 shows the coefficient estimates. Again, the Hansen-Sargan test is in favor of the instrument validity and the regularity conditions cannot be rejected. Adjustment speeds are comparable with the previous findings: 71% and 79%, respectively for Latin America and Canada. A variability tests on these two inputs yields $\chi^2(6) = 37.60$ with a probability value of 0.000, which decidedly reject the hypothesis.

Elasticity Estimates

In the received literature, the multinational responsiveness to relative price variations is investigated by estimating price (own-price and cross-price) elasticities. In this respect, our intertemporal analysis can produce novel insights as we can distinguish between short run and long run behavior. Nonetheless, a problem seems to emerge here. That is, the estimation of the long run stock price elasticities requires estimating the steady state levels $\bar{l}(p_l, y) = -(r - A_{p_l l})^{-1} [A_{p_l l} A^i_{p_l} + ry A_{p_l l} (A_{p_l y} + A_{p_l p_l} p_l)]$. Unfortunately, the latter depend on the sector specific effect $A_{p_l l} A^i_{p_l}$, which is not consistently estimated when T is small. We work around this problem by dwelling upon the long run flow elasticities, which measure the response of the annual change in the steady state employment to a relative price variation. Differently from the long run stock elasticities, flow elasticities are identified under our parametrization as the derivatives with respect to the prices are the same for both stocks and flows and the latter do not depend on the sector specific effect.

For computational ease, we get the price elasticity estimates from the two more parsimonious specification of Models 2 and 3. Tables 5 and 6 present the values of the medians of short-run stock and flow and long-run flow price elasticities for this two models. We calculate both the elasticities and the corresponding standard errors at each sample observation. The elasticities that are more frequently significant are parents and Latin American affiliates own price elasticities and cross price elasticities between parents and Latin America and of employment in Europe in response of a wage change in Latin America. The short-run and long-run elasticities present the same sign in both specifications, and, notably, in most of the cases the elasticities' values are not very different.

Taking previous indications about elasticities' significance into account, we can now analyze the complementarity/substitution issue. Tables 5 and 6 show that, in the short run, there is substitution between affiliates located in Latin America and parents in the US. This is no longer true in the long run, where the relationship is of complementarity. At the same time, an increase in the relative wage of affiliates located in Latin America causes an increase in employment in Europe, both in the short and long run. How to interpret this puzzle? A possible story is as follows. An increase in the relative wage of affiliates located in Latin America reduces the desired stock of labor in this area. This increases the gap between actual and desired employment in Latin America making adjustment costs higher. As a consequence, the multinational rises employment in the parent to maintain its output level. However, the optimal response to the changing convenience of producing in Latin America compared to other geographical areas, will be to modify more deeply the location of production within the multinational. A more formal discussion of the substitutability/complementarity reversion is offered in Appendix B, analysing the issue within the context of a slight extension of the example in Section 2.

Long-run cross price elasticities show that the response to a wage increase in Latin American affiliates will be an expansion of employment in Europe and a contraction of the workforce in the US and in Latin America. Because of the presence of high transport costs and trade barriers, the increase of Latin American wage, changing the relative convenience to produce in the Western Hemisphere and export to Europe, leads to a delocation of activities towards Europe. Proximity to the final market seems to be in this case one of the most important determinant of location decision.

The result of complementarity between employment at home and in Latin American affiliates in the long run, suggests a vertical division of activities among countries with different workforce skill levels.

A complementarity relationship between employment in Latin American affiliates and in the US parents seems to emerge also from the long-run cross price elasticities with respect to Europe wage changes. While the immediate response to an increase of European wage is a contraction of parent employment and a rise in Latin American labor demand to take advantage of wage differentials, in the long-run, the sign of parent elasticities becomes positive. It seems plausible that in the light of growing costs of performing activities in developed countries, the multinational can decide to increase the delocation of labor intensive stages of production towards developing regions and maintain the skilled intensive activities in the parent. In the long-run, due to the vertical link between different stages of production, the expansion of activities performed by Latin American affiliates may tend to bring with it an expansion of activities performed by the parents.

Conclusions

In this study we have developed a dynamic model of the multinational firm that explicitly allows for the presence of input adjustment costs. Building upon such a model, we have then carried out an empirical application to industry level data on US multinationals.

Treating labor in different geographical locations as separate factors of production, we test for the presence of adjustment costs and also estimate short-run and long-run cross-price elasticities. We find evidence of significant adjustment costs for employment in Latin American and Canadian affiliates. We also find that, in some instances, due to the presence of slow input adjustments, the complementarity/substitution relationship between employment in different international locations is reversed from the short to the long-run.

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Appendix A

For the sake of completeness, we spell out the procedure devised in Bruno (1999), generalizing the Arellano and Bond (1991) GMM estimator to the case of a system of equations with cross-equation parameter restrictions.

Assume that we have a panel of N individuals and that each individual i is observed for $T_i \geq 3$ periods of time. Consider a system of M equations with random disturbances $\varepsilon = \left[\varepsilon^{1\prime}, \ldots, \varepsilon^{M\prime}\right]^{\prime}$ which have the properties defined in Section?. We wish to estimate a vector of parameters θ, assuming that the dimension of θ guarantees the identification. Let Z be the $\sum_{i=1}^{N}(T_i - 2) \times k$ matrix of instruments as derived in Section ? For each equation, define the corresponding $\sum_{i=1}^{N}(T_i - 2) \times 1$ vector of orthogonality conditions $m^i(\theta) = Z^{\prime}(\Delta\varepsilon^i)$, $i = 1, \ldots, M$, and then stack them in a single vector $m(\theta) = [m^{1\prime}(\theta), \ldots, m^{M\prime}(\theta)]^{\prime}$. Let A_N be the $M \sum_{i=1}^{N}(T_i - 2) \times M \sum_{i=1}^{N}(T_i - 2)$ optimal weighting matrix. The optimal GMM estimator is then given by

$$\hat{\theta} = \arg\min_{\theta}\left[m^{\prime}(\theta)\ A_N\ m(\theta)\right].$$

Hansen (1982) showed that the optimal choice for A_N is $\left(\frac{1}{N}\,\widehat{m}\,\widehat{m}^{\prime}\right)^{-1}$, where $\widehat{m}$ is an estimate of $m(\theta)$ from a consistent first step estimator $\hat{\theta}_1$. As a result, a feasible GMM estimator can be obtained in two steps. The first step estimator is obtained by slightly generalizing the Arellano and Bond procedure. Let H_i be a $(T_i - 2)$ square matrix that has twos in the main diagonal, minus one in the first subdiagonals and zeroes otherwise. Then, form a $\sum_{i=1}^{N}(T_i - 2)$ block diagonal matrix H, with H_i = block i of H. Let $\bar{A}_N = \left(\frac{1}{N}\,Z^{\prime}\,H\,Z\right)^{-1}$ and, finally, get the first step estimator

$$\hat{\theta}_1 = \arg\min_{\theta}\left[m^{\prime}(\theta)\ \bar{A}_N\ m(\theta)\right].$$

Now, use $\hat{\theta}_1$ to construct $\widehat{m}$, which yields the second step estimator

$$\hat{\theta}_2 = \arg\min_{\theta}\left[m^{\prime}(\theta)\ \left(\frac{1}{N}\,\widehat{m}\,\widehat{m}^{\prime}\right)^{-1}\ m(\theta)\right].$$

$\hat{\theta}_1$ and $\hat{\theta}_2$ are asymptotically equivalent if the ε_{it}^{s}, $s = 1, \ldots, M$, $i = 1, \ldots, N$, $t = 3, \ldots, T_i$ are independent and homoskedastic across individuals, equations and over time.

Appendix B

We address the issue of the sign reversion from the short run to the long-run within a slight generalization of the example in Section 2. The production process now involves two quasi-fixed inputs, employment at region *A* and employment at region *B*, and a perfectly variable input, employment at *C*. The multinational company chooses the employment paths in the three regions to minimize its discounted flow of costs over an infinite horizon, as in (ref: minval). We assume that the intertemporal optimization problem yields the following employment demand functions at regions *A* and *B*

$$l^A = L_A - K_{AA}\widetilde{w}_A + K_{AB}\widetilde{w}_B + J_{AA}l^A_{-1} + J_{AB}l^B_{-1}$$
$$l^B = L_B - K_{BB}\widetilde{w}_B + K_{AB}\widetilde{w}_A + J_{BB}l^B_{-1}$$

where $\widetilde{w}_i$ denotes the wage paid to workers at region $i = A, B$ normalized by the wage paid to workers at C . The reader can notice that the optimal employment profiles are such that the employment adjustment in region *B* affects the adjustment in region *A*, with the converse not true. Then, assume that the matrix $\{J_{ij}\}$ is stable and that the steady states $\bar{l}^B = (L_B - K_{BB}\widetilde{w}_B + K_{AB}\widetilde{w}_A)/(1 - J_{BB})$ and $\bar{l}^A = (L_A - K_{AA}\widetilde{w}_A + K_{AB}\widetilde{w}_B + J_{AB}\bar{l}^B)/(1 - J_{AA})$ are always non-negative.

Now, the short-run cross-price elasticities are

$$a^{s-r}_{AB} = K_{AB}\frac{l^A}{\widetilde{w}_B}$$
$$a^{s-r}_{BA} = K_{AB}\frac{l^B}{\widetilde{w}_A},$$

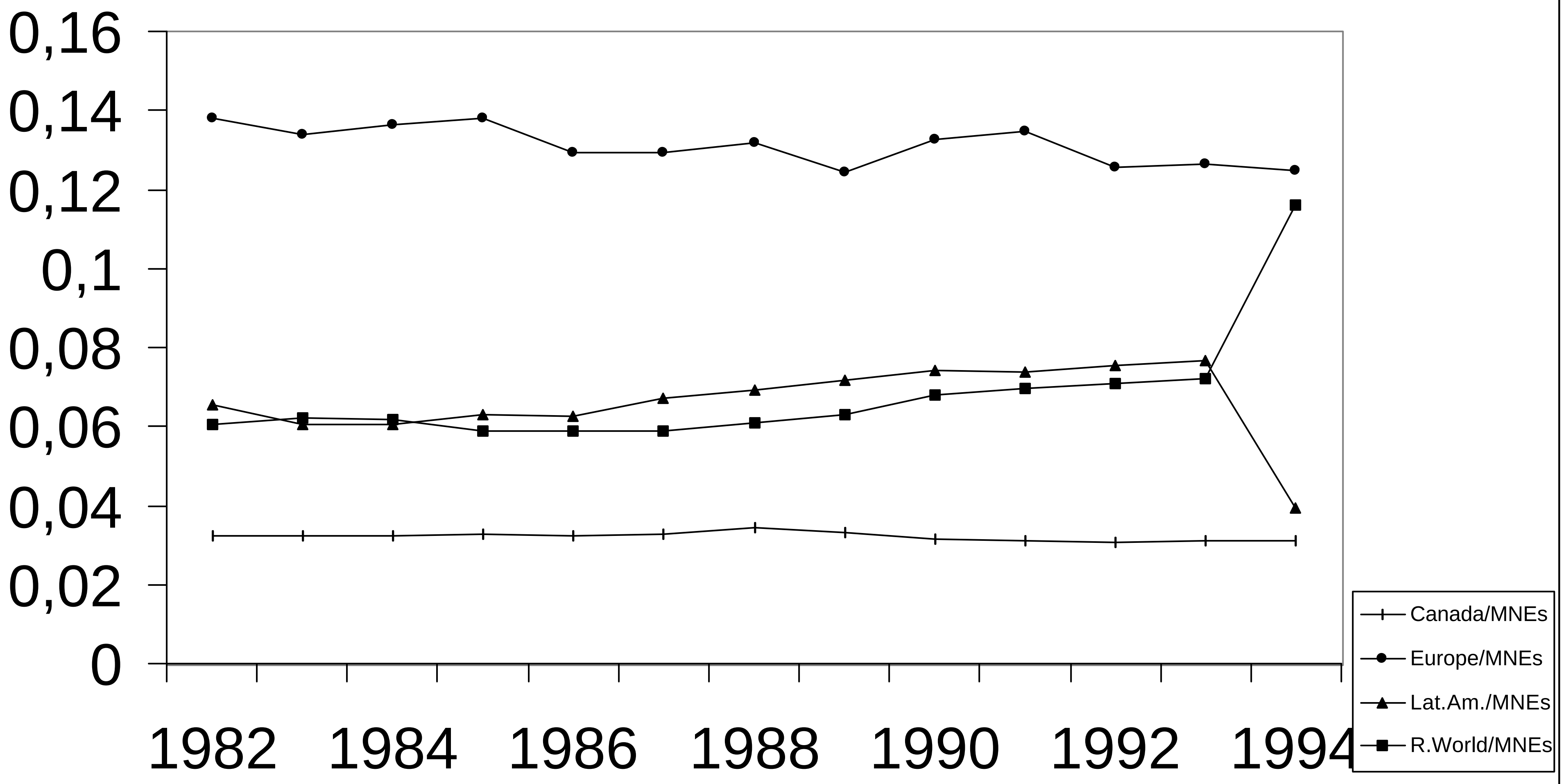
while the long-run cross-price elasticities are

$$a^{l-r}_{AB} = \frac{K_{AB}(1 - J_{BB}) - J_{AB}K_{BB}}{(1 - J_{BB})(1 - J_{AA})}\frac{\bar{l}^A}{\widetilde{w}_B}$$
$$a^{l-r}_{BA} = \frac{K_{AB}}{(1 - J_{BB})}\frac{\bar{l}^B}{\widetilde{w}_A}.$$

It is not hard to see that the short-run response of l^A to a change of $\widetilde{w}_B$, a^{s-r}_{AB}, does not need have the same sign as the corresponding long-run response, a^{l-r}_{AB}, since no sign restriction applies on $K_{AB}(1 - J_{BB}) - J_{AB}K_{BB}$. This is not the case for l^B as the adjustment at region *A* has no impact on the adjustment at *B*.


Figure 1 - Affiliate Employment Share
Employment Share
0,315
0,31
0,305
0,3
0,295
0,29
0,285
0,28
0,275
0,27
0,265
1982
1985
1988
1991
1994
Affiliates Employment/Multinationals employment

Figure 2 - Affiliate Employment Shares (Geographical locations of affiliates)



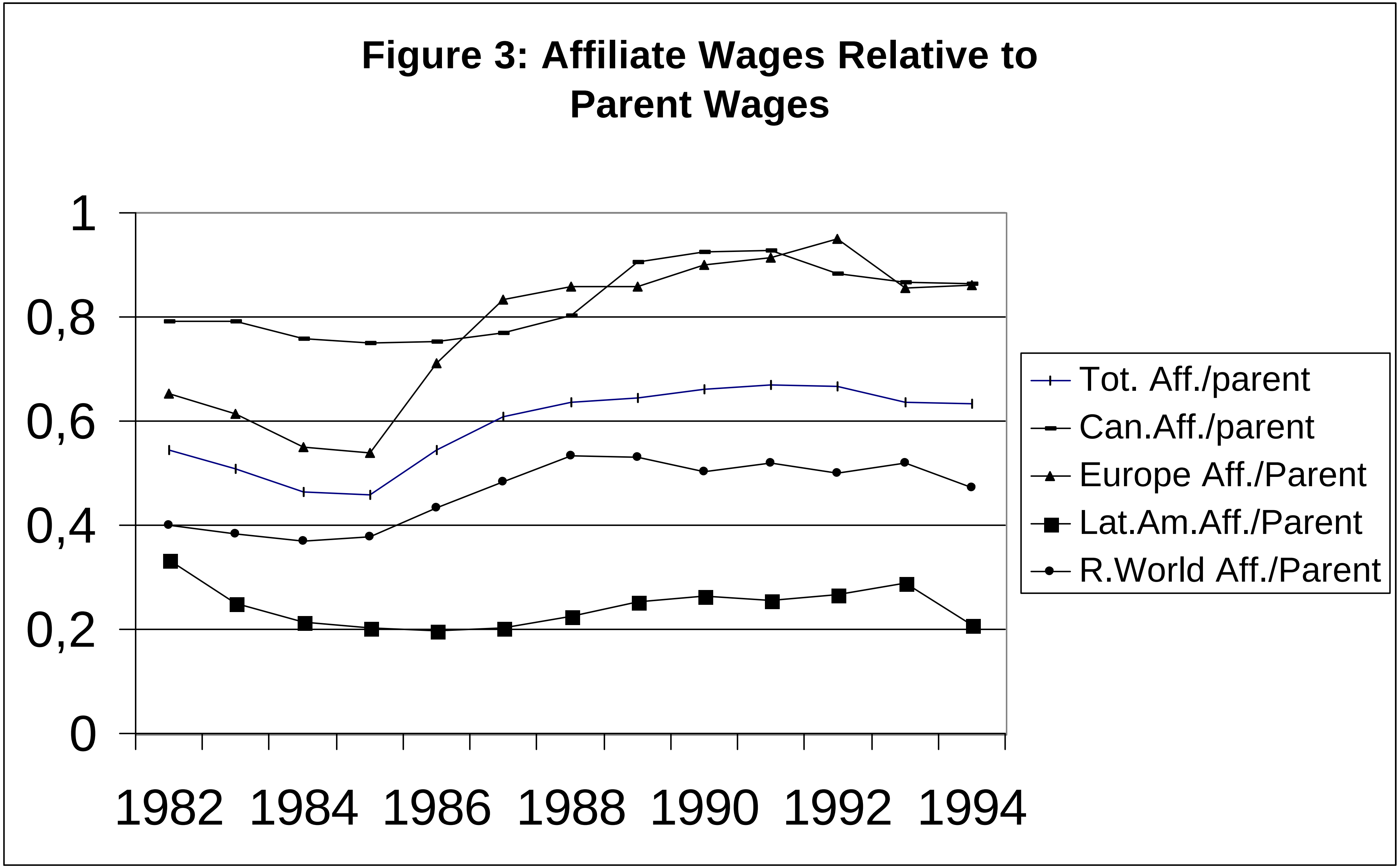


Figure 3: Affiliate Wages Relative to Parent Wages
1
0,8
0,6
0,4
0,2
0
1982
1984
1986
1988
1990
1992
1994
Tot. Aff./parent
Can.Aff./parent
Europe Aff./Parent
Lat.Am.Aff./Parent
R.World Aff./Parent

Table 1 - Autocorrelation coefficients of employment annual changes

	Autocorrelation coefficients
Parents	0.05
Europe	-0.04
Canada	0.11
Latin America	0.17
Rest of the World	0.05

Source: our calculations on BEA data retrieved from Feenstra (1997)

Table 2 - Parameters estimates from Model 1

Test of overidentifying restrictions = 169.728 [.798]
Degrees of freedom = 186
Number of Observations = 100

Parameter	Estimate	Standard Error	t-statistic	P-value
BPL	.568238E-04	.912487E-03	.062274	[.950]
BEL	.454541E-03	.928995E-03	.489282	[.625]
BCL	.110171E-02	.736590E-03	1.49569	[.135]
BLL	-.798228E-02	.277707E-02	-2.87436	[.004]
BL	-.435562E-02	.329391E-02	-1.32232	[.186]
BPC	-.659348E-03	.426136E-03	-1.54727	[.122]
BEC	.197868E-05	.430849E-03	.459252E-02	[.996]
BCC	.844445E-04	.381751E-03	.221203	[.825]
BC	.356319E-03	.163229E-02	.218294	[.827]
BPE	.525571E-03	.760810E-03	.690804	[.490]
BEE	.162589E-03	.819955E-03	.198290	[.843]
BE	-.284898E-02	.292148E-02	-.975183	[.329]
BPP	-.012847	.403238E-02	-3.18604	[.001]
BP	.106808	.017316	6.16808	[.000]
APP	1.16195	.065115	17.8447	[.000]
APE	.482769	.900293	.536236	[.592]
APC	4.45659	1.60969	2.76860	[.006]
APL	-.750719	.918484	-.817346	[.414]
AEE	.995513	.125123	7.95628	[.000]
AEP	-.907176E-02	.902607E-02	-1.00506	[.315]
AEC	.085477	.247283	.345663	[.730]
AEL	-.275293	.132871	-2.07188	[.038]
ACC	.850586	.128180	6.63589	[.000]
ACP	-.487665E-02	.472335E-02	-1.03245	[.302]
ACE	-.048261	.065741	-.734117	[.463]
ACL	-.026677	.068640	-.388649	[.698]
ALL	.766117	.125560	6.10158	[.000]
ALP	-.809472E-02	.810779E-02	-.998388	[.318]
ALE	-.024391	.112389	-.217021	[.828]
ALC	-.332180	.232545	-1.42846	[.153]

Equation PLEQ Dependent variable: PL
R-squared = .987

Equation ELEQ Dependent variable: EL
R-squared = .997

Equation CLEQ Dependent variable: CL
R-squared = .988

Equation LLEQ Dependent variable: LL
R-squared = .986

Table 3 - Parameter estimates from Model 2

Test of overidentifying restrictions = 164.105 [.999]
Degrees of freedom = 226
Number of Observations = 100

Parameter	Estimate	Standard Error	t-statistic	P-value
BEL	.594694E-03	.756043E-03	.786587	[.432]
BLL	-.545777E-02	.189902E-02	-2.87400	[.004]
BL	-.377688E-02	.250383E-02	-1.50844	[.131]
BPC	-.369504E-03	.289529E-03	-1.27622	[.202]
BCC	.126934E-03	.289763E-03	.438061	[.661]
BC	.165300E-02	.114075E-02	1.44905	[.147]
BPE	.684993E-03	.570988E-03	1.19966	[.230]
BEE	-.413138E-03	.592100E-03	-.697749	[.485]
BE	-.269494E-02	.199932E-02	-1.34792	[.178]
BPP	-.013168	.322122E-02	-4.08789	[.000]
BP	.099519	.013709	7.25967	[.000]
APP	1.13588	.060174	18.8764	[.000]
APC	4.09179	1.61589	2.53222	[.011]
APL	-1.42301	.720949	-1.97380	[.048]
AEE	.921465	.099056	9.30242	[.000]
AEL	-.257099	.110515	-2.32636	[.020]
ACC	.846345	.107573	7.86765	[.000]
ACP	-.553037E-02	.398099E-02	-1.38920	[.165]
ALL	.835258	.101020	8.26826	[.000]
ALP	-.939341E-02	.717473E-02	-1.30924	[.190]
ALC	-.450059	.192459	-2.33846	[.019]

Equation PLEQ Dependent variable: PL
R-squared = .986

Equation ELEQ Dependent variable: EL
R-squared = .997

Equation CLEQ Dependent variable: CL
R-squared = .989

Equation LLEQ Dependent variable: LL
R-squared = .986

Table 4 - Parameter estimates from Model 3

Test of overidentifying restrictions = 173.185 [.998]
Degrees of freedom = 230
Number of Observations = 100

Parameter	Estimate	Standard Error	t-statistic	P-value
BEL	.891905E-03	.732037E-03	1.21839	[.223]
BLL	-.637844E-02	.187239E-02	-3.40657	[.001]
BL	-.484929E-02	.259758E-02	-1.86685	[.062]
BPC	-.321196E-03	.281476E-03	-1.14111	[.254]
BCC	.153834E-03	.286047E-03	.537792	[.591]
BC	.120974E-02	.104726E-02	1.15515	[.248]
BPE	.383369E-03	.423633E-03	.904954	[.365]
BEE	-.337907E-03	.470017E-03	-.718924	[.472]
BE	-.283454E-02	.168699E-02	-1.68023	[.093]
BPP	-.014150	.336933E-02	-4.19955	[.000]
BP	.105089	.013685	7.67935	[.000]
APC	4.11989	1.61991	2.54328	[.011]
APL	-1.26049	.698548	-1.80444	[.071]
AEL	-.302182	.098167	-3.07826	[.002]
ACC	.860225	.105782	8.13202	[.000]
ALL	.781014	.099463	7.85235	[.000]
ALC	-.443708	.199031	-2.22934	[.026]

Equation PLEQ Dependent variable: PL
R-squared = .986

Equation ELEQ Dependent variable: EL
R-squared = .998

Equation CLEQ Dependent variable: CL
R-squared = .989

Equation LLEQ Dependent variable: LL
R-squared = .985

Table 5 - Stock and flow cross price elasticities (Model Two) - MEDIANS

Price	Parent	Europe	Canada	Latin America
Short-run Stock				
W - Parent	-0,45209*	0,0949	-0,12909	0,065543*
W - Europe	-0,001306	-0,053393	-0,001524	0,077411
W - Canada	0,002304	0,0000	0,046897	-0,009607
W - Latin America	0,047012*	0,06687*	0.0000	-0,23128*
Short-run Flow				
W - Parent	-5,46690*	3,35760*	-3,03428*	1,12283*
W - Europe	-0,014377	-1,95163	-0,032722	1,36463
W - Canada	0,022758	0,0000	1,11093	-0,16745
W - Latin America	0,51559*	2,31217*	0,0000	-4,33215*
Long-run Flow				
W - Parent	-4,54911*	3,56680*	-5,64566*	-0,084430*
W - Europe	0,18838	-1,38424	0,0083489	1,35341
W - Canada	-0,11270	0,0065113	1,67101	0,012698
W - Latin America	-0,053042*	0,48155	-0,015544	-3,47157*

Note: * indicates significance at the 10 percent level.

Table 6 - Stock and flow cross price elasticities (Model Three) - MEDIANS

Price	Parent	Europe	Canada	Latin America
Short-run Stock				
W - Parent	-0,45538*	0,061071	-0,1502	0,032089
W - Europe	-0,013587	-0,062905	0.0000	0,11059
W - Canada	0,006727	0.0000	0,055307	-0,012099
W - Latin America	0,049546*	0,098683*	0.0000	-0,25532*
Short-run Flow				
W - Parent	-5,40272*	2,16991	-3,75306*	0,56312
W - Europe	-0,15247	-2,35175	0,0000	2,01460
W - Canada	0,072329	0,0000	1,38332	-0,22039
W - Latin America	0,56647*	3,15450*	0,0000	-4,55033*
Long-run Flow				
W - Parent	-4,72664*	2,04044*	-4,29439*	-0,042163
W - Europe	0,11571	-1,32898	0,0000	1,68733*
W - Canada	-0,096054	0,010291	1,75029	0,016168
W - Latin America	-0,058824*	0,87913	0,0000	-3,40103*

Note: * indicates significance at the 10 percent level.