Are Regional Trading Partners “Natural”?

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A central statement of the theory of natural trading partners is that preferential trading with regional trading partners is less likely to be trade diverting and therefore geographically proximate partners are to be considered “natural” partners for preferential arrangements. This paper examines this question empirically. The analytical framework involves a general equilibrium model of preferential trade and an econometric model with tight links to this theory. This framework is used to implement tests of the natural trading partners hypothesis using U.S. trade data for the years 1964–95: Welfare changes that would result from preferential tariff reductions by the United States against various trading partners are first estimated, and correlations with bilateral “distance” measures (with and without controls for income levels) are then examined. Since the argument for “natural” trading partners is based on the greater likelihood of geographically proximate countries to be more *significant* trading partners, correlations between the welfare change estimates and bilateral trade volume are examined as well. Both geographic proximity and trade volume are found to have no effect. Thus this paper is unable to find any support for the natural trading partners theory in U.S. data.

I am extremely grateful to Lars Hansen, Tom MaCurdy, Arvind Panagariya, Alan Winters, and an anonymous referee for many helpful suggestions on methodology and exposition. I am also grateful to seminar participants at Brown University; University of California at Berkeley, Irvine, Los Angeles, San Diego, and Santa Cruz; Florida International University; Northwestern University; Stanford University; and the 1997 National Bureau of Economic Research Summer Institute for helpful comments and suggestions; to Joseph Hebr at U.N. Statistics Division and Patricia Burke at U.S. Customs for their help in putting the data together; and to Sajjid Chinoy, Mike Robbins, Vicharee Vichit-Vadakan, and Narayanan Subramanian for superb research assistance. Financial assistance from Brown University in the form of a Salomon faculty research grant is gratefully acknowledged, as is the hospitality and the intellectual and financial support of the Center for Research on Economic Development and Policy Reform at Stanford University, where part of this research was conducted.

[*Journal of Political Economy,* 2003, vol. 111, no. 1]

I. Introduction

The question of preferential arrangements is a long-standing one. Indeed, this issue has been the subject of active theoretical and empirical research and debate ever since the signing of the General Agreement on Trade and Tariffs (GATT), whose Article XXIV, in a significant exception to the GATT's own central principle of nondiscrimination, expressly permitted the formation of preferential trading arrangements (PTAs) in the form of free-trade areas (FTAs) and customs unions between its member countries.

That economists have been divided on the wisdom of such arrangements is not surprising; it was precisely in the context of PTAs that the complexities of second-best were first discovered. Specifically, as Viner (1950) first established in his classic analysis, PTAs, in contrast to unilateral (nondiscriminatory) trade liberalization, give rise to both trade creation and trade diversion, whose net effect on the welfare of member countries and the rest of the world is, in general, ambiguous. This ambiguity led economists to subsequently refine the theory in an attempt to determine member (or partner) country characteristics that would ensure welfare improvement with PTAs. Nevertheless, as Panagariya (1997) notes, much of the early research on this topic, pioneered by Viner (1950), Meade (1955), Lipsey (1960), and Johnson (1962) and later somewhat synthesized by McMillan and McCann (1981), yielded results that were generally considered to be mostly taxonomic and to have limited practical applicability and operational significance.

In this context, and reflecting the renewed interest in the literature on issues relating to economic geography, three recent and influential papers (Wonnacott and Lutz 1989; Krugman 1991; Summers 1991) have suggested a criterion that is remarkably simple and whose application would perhaps require only the most readily available data: geographic proximity. As Bhagwati (1993) has pointed out, the idea discussed in these papers is that we should encourage countries to enter into preferential arrangements with geographically proximate countries rather than with distant ones, because the former would more likely be trade creating, lead to a larger improvement in welfare (of the home country), and thus be "natural," and the latter would more likely be trade diverting.[1] As Bhagwati further notes, this argument itself rests on a syllogism, whose first premise is that geographically proximate countries

[1] Thus, e.g., Wonnacott and Lutz (1989) state that "trade creation is likely to be great, and trade diversion small, if the prospective members of an FTA are natural trading partners. Several points are relevant: Are the prospective members already major trading partners? If so the FTA will be reinforcing …. Are the prospective members close geographically? [Preferential] groupings of distant nations may be inefficient."

have higher volumes of trade with each other than more distant ones do and whose second is that trade blocs between countries that already trade disproportionately are less likely to divert trade.[2]

In addition to the academic interest in this idea, the question of natural trading partners is immensely interesting for policy reasons. Many existing preferential trading arrangements are indeed regional. In addition, many extensions of existing arrangements along regional lines—such as the expansion of the North American Free Trade Agreement (NAFTA) to include Chile, Argentina, and other South American countries or that of the European Union to include countries from eastern and central Europe—are currently being debated and discussed in policy circles.

Is it the case that welfare improvement is greater when liberalizing tariffs preferentially with respect to geographically proximate partners? With respect to significant trading partners? Do the data support the idea of natural trading partners? These are the questions that this paper attempts to investigate empirically.

To be sure, there already exist a number of econometric studies that examine the effects of *specific* trade blocs (such as the European Community). None, however, has addressed the questions that this paper attempts to tackle.[3] Also, as Srinivasan, Whalley, and Wooton (1993) have pointed out in their comprehensive survey on measuring the effects of regionalism, many of these "ex post" econometric studies of specific preferential arrangements, which mostly involve econometric analysis of intra- and inter-bloc trade patterns, are largely unreliable because of problems such as misspecification and simultaneity bias and parameter value instability. Finally, and most important, given the present context, a significant problem with these studies is that they usually lack microeconomic underpinnings, which makes the *welfare* analysis of even *actual* arrangements difficult and precludes entirely the possibility of welfare comparisons of alternate *potential* PTAs.

[2] Thus Krugman (1991) states that "to reemphasize why this matters: if a disproportionate share of world trade would take place between trading blocs even in the absence of a preferential trading agreement, then gains from trade creation within blocs are likely to outweigh any possible losses from external trade diversion."

[3] A partial exception is the recent and well-known work of Frankel (1996), which investigates trade patterns between countries using the "gravity" framework and finds that "distance" emerges as being a significant determinant of trade flows (after income levels have been conditioned on)—thus providing support for the idea that in a broad cross section of countries, geographical proximity matters in determining trade flows. However, the gravity framework itself does not (and indeed cannot directly) provide welfare estimates and thus cannot address the issue that this paper attempts to address regarding the merits of a policy of regionalism on welfare grounds. Other features of this approach also render it unsuitable for addressing the question at hand. See, e.g., the criticisms of Hummels (1998) and Srinivasan (1998). On the theories underlying the success of the gravity equation itself, see the recent paper by Evenett and Keller (2002).

In contrast, I begin here by outlining a simple general equilibrium framework (detailed in Sec. II) that captures the essential elements of the second-best problem of preferential tariff reduction. This framework helps identify key parameters, which are then required to be estimated to get to estimates of the overall welfare effect of preferential tariff reductions in favor of any partner country.[4] The econometric framework used here is a version of the well-known "Rotterdam model" developed by Theil (1965) and Barten (1967) (familiar from its extensive use in applied demand analysis in earlier decades and also from its innovative and more recent application in estimating trade elasticities by Marquez [1994]) and has the benefit that it easily permits welfare analysis and comparisons since, by design, it is firmly grounded in an optimization framework. In the actual implementation here, disaggregated U.S. trade data from the years 1965–95 are used to estimate the potential welfare effects from trade creation and trade diversion that would result from a preferential reduction in U.S. tariffs with respect to imports from various geographically dispersed potential partner countries. To test the natural trading partners idea, these estimates of the overall welfare effects are correlated with the distance between the United States and the corresponding partner countries (with and without conditioning for income levels of the partner countries) and with bilateral trade volume. The main results are as follows: First, estimates of trade creation and trade diversion associated with preferential liberalization of U.S. tariffs against various partner countries are obtained. Second, the correlation between the overall welfare effect and distance is found to be statistically insignificant, and therefore the null that "distance does not matter" cannot be rejected. Nor can the null that the "bilateral volume of trade does not matter." Thus, in U.S. data, I am unable to find any evidentiary support for the natural trading partners theory.

Overall then, this paper makes two contributions. The first is a methodological one: This paper brings together and uses theory and an estimation framework that have far tighter links to each other and impose far fewer restrictions on functional forms than is traditional in the literature on preferential trade agreements. Second, in the actual implementation using U.S. data, it obtains and contrasts welfare estimates from preferential tariff reduction by the United States against partner countries at various levels of geographic proximity, thereby contributing to the ongoing policy debate concerning the benefits of regionalism in preferential trading. It should be emphasized that the findings relate

[4] Rather than examine the welfare effects of trade blocs as such, this paper actually looks at welfare effects of unilateral preferential tariff reduction instead—an approach that is entirely consistent with the Vinerian frameworks in which the discussion of natural trading partners has often been set (as, e.g., in Krugman [1991] and Bhagwati and Panagariya [1996]).

to U.S. trade data alone. Analysis of data for other countries may well find that geographically proximate countries make for better preferential trading partners in some instances. The present analysis, nevertheless, points to the difficulty of identifying systematic economic criteria in choosing partners for preferential trading: Outcomes may simply be highly sensitive to context.

The rest of this paper is structured as follows: Section II presents the basic model. Section III discusses the econometric methodology. Section IV describes the data. Section V discusses the estimation results. Section VI presents conclusions.

II. The Model

In classic Vinerian fashion, consider a trading world that is composed of three countries: country A, its prospective partner country B, and a third country C, representing the rest of the world. Each country produces only a single good, some of which it exports to pay for its consumption (imports) of the other two goods. When the border price of each good is normalized to be one,[5] country A's budget constraint, representing the equality between expenditures and total revenues (revenues from sales of any goods produced plus import tax revenue), can be expressed as

$$E(1, 1 + t_B, 1 + t_C, W) = R(1, 1 + t_B, 1 + t_C, \bar{V}) + t_B M_B + t_C M_C, \quad (1)$$

where E is the expenditure function associated with country A, R is the revenue function (i.e., revenue derived from sales of any goods produced), W denotes country A's welfare, $\bar{V}$ denotes the (fixed) factor supplies used in production in A, and t_B, t_C, M_B, and M_C denote tariffs imposed against and imports from countries B and C, respectively. Equation (1) represents the budget constraint and the expenditure and revenue functions in the general form that is traditional in trade theory. Note first that the (Armington) assumption that each country produces only a single good and that factor supplies are fixed implies that the partials of the revenue function with respect to prices are zero. To get to the effect of a preferential reduction in tariffs imposed by country A against country B (holding tariffs against C fixed at $t_C = \bar{t}_C$), let us totally differentiate (1) and let E_i denote the partial derivatives of E with respect to the ith domestic price to obtain

$$E_B dt_B + E_W d_W = t_B dM_B + M_B dt_B + \bar{t}_c dM_C. \quad (2)$$

[5] Thus the small-country assumption is made here and terms of trade effects are ignored. The theoretical analysis can, of course, be readily extended to allow for terms of trade changes.

Since the partials of the expenditure function, E_i, denote consumption of the ith good, it follows that $E_B = M_B$, and (2) therefore reduces to

$$E_W dW = t_B dM_B + \bar{t}_C dM_C, \tag{3}$$

where $E_W > 0$, since it is simply the inverse of the marginal utility of income (which helps convert the real income changes on the right-hand side into welfare units). Expression (3) has the familiar intuitive interpretation: For welfare improvement to be guaranteed, imports from both the partner country and the rest of the world should increase. If, alternatively, imports from the partner country increase, $dM_B > 0$, implying trade creation, but imports from the rest of the world decrease, $dM_C < 0$, implying classic trade diversion, welfare might drop instead.

To relate equation (3) above to country characteristics, we can make use of the fact that the compensated import demand functions, M_B and M_C, themselves are a function of prices and welfare. Thus they can be expressed as

$$M_B = M_B(1,\ 1 + t_B,\ 1 + \bar{t}_C,\ W) \tag{4}$$

and

$$M_C = M_C(1,\ 1 + t_B,\ 1 + \bar{t}_C,\ W). \tag{5}$$

Totally differentiating (4) and (5) gives us

$$dM_B = M_{BB} dt_B + M_{BW} dW \tag{6}$$

and

$$dM_C = M_{CB} dt_B + M_{CW} dW. \tag{7}$$

Substituting (6) and (7) into (3) gives us

$$(E_W - t_B M_{BW} - \bar{t}_C M_{CW}) dW = (t_B M_{BB} + \bar{t}_C M_{CB}) dt_B. \tag{8}$$

Since E is homogeneous of degree one in prices, E_W is also homogeneous of degree one in prices. Using Euler's theorem, we then have

$$E_W = E_{AW} + (1 + t_B) E_{BW} + (1 + t_C) E_{CW}. \tag{9}$$

Substituting (9) into (8) gives us a final expression for welfare similar to the one derived by McMillan and McCann (1981):

$$H dW = (t_B M_{BB} + t_C M_{CB}) dt_B, \tag{10}$$

where, from (8) and (9), H is given by

$$H = E_{AW} + E_{BW} + E_{CW} = \left[\frac{\partial (E_A + E_B + E_C)}{\partial I}\right] E_W, \tag{11}$$

with I denoting income (which is equal to expenditure E). Equations (10) and (11) may be understood as follows. In (11), the partials $\partial E_i/\partial I$ denote the extent to which consumption of good i increases with a unit increase in income (with prices held fixed). Given that E_W is simply the inverse of the marginal utility of income, H then denotes the total increase in the real value of the consumption basket (i.e., the value of the consumption basket at world prices) corresponding to a unit increase in welfare. As Dixit and Norman (1980) note, HdW then simply represents real income change. One can therefore interpret the expression on the right-hand side of (10) as the real income change associated with the tariff reduction dt_B, with H serving to translate welfare changes into real income changes. Note that if all goods are normal in consumption (as will be assumed here), then $\partial E_i/\partial I > 0$ for all i, implying that H is positive as well.

Expression (10) tells us that welfare improvement is *guaranteed* if imports from the partner country are substitutes for home country output and are complementary to imports from the rest of the world. Note, however, that even if the rest of the world output and partner country output are substitutes (as will generally be the case), welfare may go up if the trade creation term on the right-hand side of (10) dominates the trade diversion term.

Two observations may be made here. First, (3) and (10) indicate that the methodology followed in many previous analyses—of simply adding up the estimated volumes of trade created and diverted (changes in the volumes of trade with respect to the partner country and the rest of the world)—is somewhat incorrect: these changes need to be weighted by the initial tariff levels as in (3) or (10). Second, given initial conditions, (11) implies that the term H is independent of the particular bilateral tariff reduction that is being considered. In other words, if we were to compare the welfare effect of a preferential reduction in tariffs by A against country B with the welfare effect when tariffs are preferentially reduced against C instead, a comparison of the right-hand side of (10) in the two cases would suffice to establish a welfare ranking. Of course, to estimate the right-hand side of (10) for preferential tariff reductions against each potential partner country, we need to estimate the own-price and cross-price effects on imports from the partner country and the rest of the world in each case. It is to this problem that this paper turns next.

III. Estimation Methodology

To estimate the own- and cross-price effects in (10), this paper uses a version of the Rotterdam model, developed by Theil (1965) and Barten (1967) and used recently in estimating U.S. trade elasticities by Marquez

(1994). The Rotterdam model embodies, by design, all the properties of utility maximization, thus recognizing the interdependence between spending decisions on domestic and foreign goods, and does not treat trade elasticities as autonomous parameters (thus avoiding the awkward problem relating to the predetermination of elasticities pointed out by Koopmans and Uzawa [1990]; see Marquez [1994] for a discussion). Individuals determine their spending on domestic and foreign goods by maximizing a utility function, $U(q_1, \ldots, q_n)$, subject to a budget constraint, $\sum_j p_j q_j = I$,[6] where I denotes income and j is a country index. Obtaining the first-order conditions for maximizing *any* $U(\cdot)$ and totally differentiating the associated system of Marshallian demands (themselves functions of income and prices) yields the following expression for the demand for the ith product:

$$(\omega_{it})d\ln q_{it} = \left[\frac{\partial(p_{it}q_{it})}{\partial I_t}\right]d\ln\left(\frac{I}{P}\right)_t + \sum_j \left[\left(\frac{p_{it}p_{jt}}{I_t}\right)\left(\frac{\partial h_{it}}{\partial p_{jt}}\right)\right]d\ln p_{jt}, \qquad (12)$$

where $\omega_{it} = p_{it}q_{it}/I_t$, $p_{jt} = (1 + \tau_{jt})p_{xjt}$ is the domestic price of good j, $d\ln p_t = \sum_j (\omega_{jt})d\ln p_{jt}$, τ_{jt} is the tariff rate on imports from j, and p_{xjt} is the border price of the jth good.

To implement (12) empirically, the Rotterdam model restricts the marginal budget share, $\mu_i = \partial(p_{it}q_{it})/\partial I_t$, and the Slutsky coefficients, $\pi_{ij} = (p_{it}p_{jt}/I_t)(\partial q_{it}/\partial p_{jt})$, to be invariant to changes in income and prices.[7] The marginal budget share measures the additional amount spent on the ith good, when income increases by one dollar. The Slutsky coefficient measures the compensated price effect of a change in the price of the jth good on purchases of the ith good. Treating these parameters as autonomous transforms (12) into the standard Rotterdam estimating equation:

$$(\omega_{it})d\ln q_{it} = \mu_i d\ln\left(\frac{I}{P}\right)_t + \sum_j (\pi_{ij})d\ln p_{jt} + r_{it}, \qquad (13)$$

where r_{it} is a random disturbance term.

Note here that estimation of the demand system (13) gives us own- and cross-price effects (the relevant π_{ij}'s) that may be substituted back into (10) in order to get estimates of welfare change due to preferential

[6] It should be pointed out that the spending decisions here, in common with much of the previous literature, suffer from at least the limitations that they ignore intertemporal substitution and that labor supply and asset-holding decisions are taken to be separable from decisions to consume domestic and foreign products.

[7] As Marquez (1994) notes, some critics of the Rotterdam model have pointed out that treating μ and π as invariant to income and prices *implies* Cobb-Douglas preferences. However, Barnett (1979) has argued convincingly that this criticism has only limited relevance for empirical work with *aggregate* data. See also Deaton and Muellbauer (1980) for a discussion.

tariff reduction. The income and price elasticities associated with (13) are μ_i/ω_{it} and π_{ij}/ω_{it}, respectively.

Note also that for the parameters of the Rotterdam system to be consistent with utility maximization, they need to satisfy the following restrictions: (1) the adding-up constraint on marginal budget shares,

$$\sum_j \mu_j = 1; \tag{14}$$

(2) homogeneity of demand,

$$\sum_j \pi_{ij} = 0 \quad \forall i; \tag{15}$$

and (3) symmetry,

$$\pi_{ij} = \pi_{ji} \quad \forall i,\ j,\ i \neq j, \tag{16}$$

which are all restrictions that are taken into account in the estimation.

In the analysis the United States is taken to be the home country and 24 different partner countries are considered. In estimating the demand system (13), one would ideally consider a full multilateral demand system that includes 26 equations (24 partner countries plus the United States plus the rest of the world). However, this leaves us with a very large number of parameters to estimate. In particular, when we are considering preferential tariff reductions against any one country, the cross-price term that we would need to include in (10) to calculate the overall welfare effect would sum over cross-price terms from the remaining 24 countries. The calculated standard errors of this sum tend to be very large—rendering virtually all the estimates of cross-price effects to be insignificantly different from zero. Since our final goal is to estimate, for each partner country, the value of the own-price effect and the cross-price effect (aggregated over the countries in the rest of the world in each case), let us proceed by aggregating the rest of the world into a single unit to begin, thereby following the approach taken in some classic studies in international trade that have estimated trade elasticities (e.g., Hickman and Lau 1973; Goldstein and Khan 1978; Geraci and Prewo 1982). Thus, for each of the 24 partner countries considered in this analysis, we can estimate a "triad" system by splitting the world up into the home country, the relevant partner country, and the rest of the world. Equation (13) then gives us a three-equation system to be estimated. In the estimation of each triad system, let us drop (as is the usual practice) one equation, the equation for demand for goods produced by the home economy, to avoid the singularity that the adding-up constraint imposes.

In estimating the demand system, we have to consider the additional issues of simultaneity bias and measurement error in the right-hand-side variables—both of which could imply a correlation between the regressors and the error terms. Simultaneity bias may arise if the home country is not "small" in the trade-theoretic sense so that a change in its tariffs on imports would result in a change in the border prices of its imports. Measurement error may arise since the prices that are included on the right-hand side are unit values rather than actual prices. We can deal with this problem of the possible correlation between the regressors and the error terms by using the method of instrumental variables. In particular, (13) is estimated jointly with four reduced-form "instrument equations" (one equation for each of the endogenous variables on the right-hand side of [13]: prices in the home country, the partner country, the rest of the world, and the real income term). The endogenous variables are specified to be functions of the exogenous variables in the system in the following manner:

$$x_j = \sum_k \beta_k X_k + \epsilon_j, \tag{17}$$

where the left-hand-side variables are the endogenous regressors in (13) and the right-hand-side variables are the exogenous variables. The exogenous variables here are growth rates of U.S. income and the growth rates of *aggregate* wage rates in the partner country, in the four major trading partners other than the partner country, and in the United States. Since trade volumes with the United States relative to domestic gross domestic product tend to be fairly small fractions, the exogeneity of the aggregate wage rate is not much of a concern. Statistical tests to test validity of the instruments used in the equations are discussed in Section V.

Under the assumption that the errors of the spending equations (13) and the reduced-form (instrumenting) equations (17) have a joint normal distribution with zero mean and constant covariance matrix, the demand system can be estimated with the method of maximum likelihood, reliance on which allows direct incorporation of the restrictions on homogeneity and symmetry associated with consumer demand theory. Likelihood ratio tests (discussed further in the next section) allow us to test these restrictions and also to test the possible exogeneity of the right-hand-side variables in (13).

Estimating (13) gives us estimates for the own- and cross-price effects in the case of each partner country. Plugging these into the right-hand side of (10) gives us an overall welfare effect—HdW—that can then be compared with corresponding values for preferential tariff reduction with respect to other countries. Comparing the welfare effect across countries, each at a different level of geographic proximity with the home country and significance as trading partners, allows us to test the natural trading partners idea.

IV. Data

For our estimation, we need *bilateral* price and quantity information on imports from trading partners as well as wage rates in all the partner countries. The present analysis employs U.S. imports data for the period 1965–95 obtained from the United Nations Statistics Division in New York. This data set provides time-series information on import values (measured in U.S. dollars, cost, insurance, and freight) and separately on quantities of trade flows at the three-digit level. Information on bilateral tariff rates is also required: the *bilateral* customs collection rate was used here as the tariff measure. Data on customs collections for the period 1990–95 were obtained directly from U.S. Customs. The rest, for the years 1964–89, were gathered from the U.S. Department of Commerce publication FT 990. Wage data were obtained from the International Labor Organization and national accounts publications for the corresponding years.

Aggregate bilateral price indices were constructed using the unit values that are implied by the U.N. data. For robustness, two aggregate price indices were used: the Fisher ideal index, recommended by Fisher (1927), and the Laspeyeres index. Aggregate quantity series were constructed by deflating aggregate trade flows with the corresponding price indices.[8] For consumption of domestically produced goods, the consumer price index was used as the price measure. U.S. purchases of domestically produced goods are measured as gross national product minus exports.

V. Estimation Results

Tables 1 and 2 report the estimates of the own- and cross-price effects of U.S. preferential tariff reduction with respect to 24 different countries obtained by maximum likelihood estimation of the demand system, with the Laspeyeres index and the Fisher ideal index, respectively, used as the price indices. The corresponding elasticities, calculated using beginning-of-period budget shares, are reported in tables 1 and 2 as well.

As can be seen from tables 1 and 2, all the estimates of the cross-price (trade diversion) effects are positive, indicating that the rest of the world's output and the partner country's output were substitutes and therefore that preferential tariff reductions would result in some

[8] To arrive at the results reported in this paper, price indices were constructed using all available price data for any trade basket and then that price was applied to the entire basket. The results appeared robust, however, to other methods of constructing price indices—such as those involving interpolation over missing values, the use of "chain indices" to take into account the changing composition of export baskets, and others.

TABLE 1
MAXIMUM LIKELIHOOD ESTIMATES OF OWN- AND CROSS-PRICE EFFECTS: LASPEYERES INDEX

Country	Cross-Price Effect: $\pi_{ij} \times 10^3$	Cross-Price Elasticity: π_{ij}/ω_i	Own-Price Effect: $\pi_{ii} \times 10^3$	Own-Price Elasticity: π_{ii}/ω_i	Welfare: $(HdW/Y) \times 10^6$	R^2 Partner
Argentina	.18	1.09	−.24	−1.44	1.48	.07
	(.05)	(.32)	(.08)	(.46)	(1.61)	
Australia	.28	.65	−.48	−1.11	5.12	.21
	(.07)	(.17)	(.08)	(.18)	(2.22)	
Belgium	1.11	1.61	−1.45	-2.11	8.73	.52
	(.23)	(.33)	(.16)	(.23)	(4.37)	
Brazil	.99	1.38	−.68	−.95	−7.94	.23
	(.16)	(.22)	(.27)	(.38)	(4.79)	
Canada	7.17	1.06	−11.20	−1.66	103.16	.57
	(1.49)	(.22)	(1.55)	(.23)	(42.25)	
Chile	.36	1.26	−.35	−1.22	−.26	.28
	(.03)	(.09)	(.04)	(.14)	(.78)	
France	.62	.72	−1.53	−1.78	23.30	.17
	(.25)	(.29)	(.19)	(.22)	(7.50)	
Germany	1.56	.83	−3.00	−1.59	36.86	.1
	(.41)	(.22)	(.29)	(.15)	(11.85)	
Honduras	.00	.00	−.09	−.88	2.17	.14
	(.03)	(.31)	(.04)	(.43)	(.87)	
Hong Kong	.12	.25	−2.98	−6.21	73.19	.15
	(.30)	(.63)	(.17)	(.35)	(7.91)	
Indonesia	2.40	10.42	−.97	−4.21	−36.61	.27
	(.35)	(1.53)	(.27)	(1.17)	(6.07)	
Jamaica	.03	.14	−.11	−.64	2.18	.12
	(.04)	(.23)	(.02)	(.14)	(1.22)	
Japan	8.40	2.49	−17.68	−5.23	237.56	.29
	(2.10)	(.62)	(2.02)	(.60)	(65.93)	
Korea	2.10	29.37	−1.40	−19.58	−17.92	.29
	(.23)	(3.16)	(.34)	(4.76)	(6.50)	
Mexico	3.13	3.51	−4.11	−4.61	25.09	.03
	(.50)	(.56)	(.79)	(.89)	(12.38)	
New Zealand	.04	.22	−.11	−.62	1.79	.17
	(.02)	(.11)	(.03)	(.17)	(.97)	
Peru	.26	.78	−.08	−.24	−4.61	.19
	(.06)	(.17)	(.03)	(.09)	(1.48)	
Philippines	.11	.21	−.58	−1.13	12.03	.27
	(.08)	(.17)	(.06)	(.11)	(2.36)	
South Africa	.40	1.30	−.50	−1.62	2.56	.33
	(.10)	(.33)	(.08)	(.24)	(2.84)	
Switzerland	.16	.38	−.50	−1.17	8.70	.11
	(.08)	(.19)	(.07)	(.16)	(2.90)	
Taiwan	1.39	11.05	−6.20	−49.28	123.13	.49
	(.75)	(5.96)	(.08)	(.64)	(13.95)	
Thailand	.18	3.33	−.52	−9.61	8.70	.23
	(.10)	(1.85)	(.06)	(1.11)	(1.11)	
Turkey	.03	.23	−.10	−.88	1.89	.24
	(.05)	(.47)	(.01)	(.04)	(1.29)	
United Kingdom	.79	.40	−3.61	−1.84	72.19	.07
	(.46)	(.23)	(.61)	(.31)	(18.81)	

NOTE.—Figures in parentheses are standard errors. The letter *i* denotes partner country and the letter *j* denotes the rest of the world. Elasticities were calculated using beginning-of-period budget shares of the relevant partner country in each case. Thus the elasticity corresponding to the cross-price elasticity is the proportional change in imports.

TABLE 2
MAXIMUM LIKELIHOOD ESTIMATES OF OWN- AND CROSS-PRICE EFFECTS: FISHER IDEAL INDEX

Country	Cross-Price Effect: $\pi_{ij} \times 10^3$	Cross-Price Elasticity: π_{ij}/ω_i	Own-Price Effect: $\pi_{ii} \times 10^3$	Own-Price Elasticity: π_{ii}/ω_i	Welfare: $(HdW/Y) \times 10^6$	R^2 Partner
Argentina	.21	1.27	−.43	−2.59	5.63	.04
	(.06)	(.37)	(.11)	(.66)	(1.72)	
Australia	.20	.46	−.41	−.95	5.38	.07
	(.07)	(.17)	(.08)	(.18)	(2.18)	
Belgium	.40	.58	−.93	−1.35	13.57	.44
	(.20)	(.29)	(.10)	(.15)	(4.60)	
Brazil	.56	.78	−.39	−.55	−4.35	.02
	(.10)	(.15)	(.19)	(.27)	(2.93)	
Canada	6.31	.93	−15.03	−2.22	223.22	.46
	(2.14)	(.32)	(3.00)	(.44)	(48.84)	
Chile	.38	1.32	−.35	−1.22	−.77	.57
	(.03)	(.11)	(.01)	(.03)	(.80)	
France	.97	1.12	−1.73	−2.01	19.58	.39
	(.25)	(.29)	(.21)	(.24)	(7.47)	
Germany	1.39	.74	−2.59	−1.37	30.72	.88
	(.39)	(.21)	(.23)	(.12)	(18.34)	
Honduras	.06	.61	−.13	−1.34	1.82	.01
	(.26)	(2.69)	(.05)	(.51)	(6.67)	
Hong Kong	.00	.01	−2.47	−5.15	63.11	.47
	(.28)	(.58)	(.17)	(.35)	(7.91)	
Indonesia	1.20	5.21	−1.00	−4.34	−5.12	.07
	(.34)	(1.47)	(.25)	(1.09)	(5.68)	
Jamaica	.02	.14	−.11	−.64	2.20	.08
	(.04)	(.24)	(.02)	(.14)	(1.24)	
Japan	5.85	1.73	−18.63	−5.52	327.16	.43
	(2.00)	(.59)	(2.10)	(.62)	(68.83)	
Korea	1.88	26.29	−1.24	−17.34	−16.38	.06
	(.22)	(3.04)	(.35)	(4.90)	(6.49)	
Mexico	2.17	2.43	−4.60	−5.16	62.21	.15
	(.47)	(.52)	(.53)	(.59)	(14.18)	
New Zealand	.04	.22	−.14	−.78	2.56	.20
	(.02)	(.11)	(.04)	(.20)	(1.07)	
Peru	.24	.72	−.09	−.25	−3.97	.06
	(.07)	(.20)	(.05)	(.15)	(1.73)	
Philippines	.02	.04	−.48	−.93	11.80	.38
	(.08)	(.17)	(.06)	(.11)	(1.79)	
South Africa	.21	.68	−.33	−1.07	3.07	.23
	(.06)	(.20)	(.08)	(.25)	(1.22)	
Switzerland	.15	.35	−.33	−.77	4.61	.68
	(.09)	(.21)	(.03)	(.07)	(2.50)	
Taiwan	1.06	8.43	−5.81	−46.18	121.60	.07
	(.76)	(6.04)	(.53)	(4.21)	(19.90)	
Thailand	.12	2.16	−.85	−15.71	18.76	.03
	(.18)	(3.33)	(.09)	(1.57)	(3.74)	
Turkey	.03	.25	−.15	−1.29	3.05	.23
	(.04)	(.35)	(.02)	(.15)	(1.04)	
United Kingdom	.65	.33	−3.00	−1.53	60.16	.42
	(.47)	(.24)	(.64)	(.33)	(17.63)	

NOTE.—Figures in parentheses are standard errors. The letter *i* denotes partner country and the letter *j* denotes the rest of the world. Elasticities were calculated using beginning-of-period budget shares of the relevant partner country in each case. Thus the elasticity corresponding to the cross-price elasticity is the proportional change in imports.

trade diversion. All the estimates of the own-price (trade creation) effects are negative, suggesting that these FTAs can be expected to have positive trade-creating effects as well. A simple comparison of the point estimates of the own- and cross-price effects themselves indicates that in most cases (although there are a few exceptions), own-price effects dominate cross-price effects in magnitude. Thus, roughly speaking, in these cases, if initial tariffs on the partner country and the rest of the world were equal, trade creation would outweigh trade diversion around this initial equilibrium. Finally, the implied elasticities reported in tables 1 and 2 are broadly in line with elasticity estimates reported in the literature (e.g., Marquez 1994).

Tables 1 and 2 also report some goodness-of-fit statistics. As previously noted, two equations are estimated in each triad system, the equation for import demand from the partner and import demand from the rest of the world. The equation corresponding to demand for goods from the home economy is dropped because of the adding-up constraint imposed by the budget constraint. The R^2's corresponding to the demand equation representing demand for goods from the relevant partner country are each reported in tables 1 and 2. As is readily evident, there is substantial variation in the goodness of fit, with the R^2 measures varying between .88 and .01. In most cases, however, the economic variables included on the right-hand side seem to go a reasonable distance in explaining demand variations. Tables 1 and 2 do not report the R^2's corresponding to the demand for imports from the "rest of the world" since the change in the extent of fit of this equation (as partner countries are changed) is quite small: excluding any single partner country from or including it into the aggregate category of the "rest of the world" has generally little impact on the goodness of fit of the equation representing the demand for goods from the rest of the world. These R^2's were close to about .53 with Laspeyeres prices and about .57 with Fisher prices in most cases.

Additional diagnostics are presented in table 3, which lists log likelihood ratio test statistics for joint tests of homogeneity and symmetry as well as for tests of exogeneity. The former tests the restrictions implied by utility maximization (homogeneity and symmetry) under which the estimations were carried out. The test statistic in each case, then, is distributed χ^2 with three degrees of freedom, corresponding to the one cross-price restriction and the two homogeneity restrictions imposed in the two-equation system that is actually estimated (with the critical value at the 95 percent level being 7.82). As can easily be seen, the data cannot reject the restrictions associated with homogeneity and symmetry at the 95 percent level (although in one case—Peru with the Laspeyeres case—one can reject the null at the 90 percent level). To test for exogeneity of the regressions in (13) is to test the null that the correlations

TABLE 3
Likelihood Ratio Tests of Demand Restrictions and Exogeneity

	Laspeyeres		Fisher	
Country	Demand Restrictions	Exogeneity	Demand Restrictions	Exogeneity
Argentina	4.50	2.20	5.65	3.59
Australia	5.36	8.87	5.52	10.34
Belgium	.56	9.40	2.52	9.80
Brazil	4.92	11.04	2.40	3.91
Canada	3.28	12.56	2.63	11.66
Chile	6.02	12.94	5.63	11.79
France	2.96	9.90	3.89	12.00
Germany	.76	12.56	6.03	10.08
Honduras	6.09	4.77	5.70	4.68
Hong Kong	3.98	13.34	4.08	9.34
Indonesia	3.17	13.00	6.06	11.63
Jamaica	2.45	12.05	2.35	13.93
Japan	5.87	13.21	4.35	11.75
Korea	3.23	5.03	3.10	13.08
Mexico	5.86	12.14	5.74	11.45
New Zealand	1.81	6.65	5.08	7.23
Peru	7.53	4.39	6.04	9.84
Philippines	1.68	10.61	3.16	10.22
South Africa	4.68	6.26	5.87	6.18
Switzerland	3.72	11.51	3.39	12.49
Taiwan	6.22	7.70	6.75	8.62
Thailand	2.79	12.80	7.12	13.85
Turkey	5.45	14.43	5.57	13.91
United Kingdom	2.34	2.80	1.76	9.55

Note.—Figures in parentheses are standard errors. The critical value for the χ^2_3 likelihood ratio test statistic (testing the joint restrictions of homogeneity and symmetry) at the 5 percent level is 7.82 and at the 10 percent level is 6.25. The critical value for the χ^2_8 likelihood ratio test statistic (testing exogeneity) at the 5 percent level is 15.51 and at the 10 percent level is 13.36.

between the errors of (the two equations in) (13) and (the four equations in) (17), eight correlation terms in all, are jointly zero. As table 3 reports, in every case, the χ^2_8 test statistics fall below the critical level at the 95 percent level (however, for Turkey and Thailand the null can be rejected at the 90 percent level). The null that the regressors are exogenous, that is, that trade prices are independent of U.S. bilateral trade policies, cannot be rejected.[9]

Tables 1 and 2 also present welfare estimates and the associated standard errors, which were constructed using the estimates of the own- and cross-price effects and using (10). One point needs to be noted here: this expression for welfare change, that is, (10), involves domestic prices and therefore depends on the initial tariff levels. To test the natural

[9] Also, in every case, the $F_{4,23}$ test statistics, testing the validity (Bassman tests) of the instruments (see Phillips 1983, p. 488), fall below the critical value at the 95 percent level. Thus we are unable to reject the null that the instruments are valid and uncorrelated with the error terms of the demand equations.

trading partners hypothesis, however, we are interested in how welfare changes correlate with distance, ceteris paribus, and we therefore need to construct our estimates of welfare change by considering a situation in which tariffs against all countries are at some *equal* level initially. Thus we can compute these welfare effects by substituting in (10) the average (across all countries) tariff imposed by the United States in 1994 (approximately 2.6 percent) rather than using the *actual* bilateral tariff levels at this time. Given the relative precision of estimates of own- and cross-price effects as reported in tables 1 and 2, the estimates of welfare change are also overwhelmingly significant. The magnitudes seem reasonable (and in line with other estimates in the literature) as well.[10]

To examine correlations of the welfare estimates with distance, the following regression was then run:

$$-\frac{HdW_j}{Y} = \alpha + \beta_1(\text{distance})_j + \epsilon_j, \tag{18}$$

where the distance measure used was the bilateral direct line distance (measured in thousands of miles) used by Frankel, Stein, and Wei (1995).

To see how the welfare effect was correlated with distance after conditioning for the income levels of the partner countries (as suggested by proponents of the gravity approach), the following equation was estimated:

$$-\frac{HdW_j}{Y} = \alpha + \beta_1(\text{distance})_j + \beta_2(\text{income})_j + \epsilon_j, \tag{19}$$

where income levels were simply GDP levels (measured in billions of dollars).

Finally, to see how the welfare effect was correlated with the bilateral volume of trade (since it is this that drives the economic argument for

[10] A simple calculation, albeit a highly crude one, may serve to give a better sense of the magnitudes involved here and enable comparisons with other estimates obtained in the literature. Consider, e.g., the welfare estimate reported for Japan in table 1, which was itself calculated using an initial tariff level of 2.6 percent. For comparison with other estimates in the literature, where initial tariff rates are higher, we can recalculate the welfare estimate for Japan assuming initial U.S. most favored nation tariffs to be 30 percent instead, and we can undertake a 15 percent preferential tariff reduction in favor of Japan. This gives us a welfare gain for the United States of 0.026 percent of GDP. This number may appear small, and it is. Note, however, that it is obtained in a preferential tariff reduction context (where trade diversion effects damp welfare improvement) and that it results from a tariff reduction against a single country (Japan) whose share of the import basket is roughly a tenth. In consideration of this, it should be clear that this estimate is in line with other estimates of static welfare gains obtained in the literature. See, e.g., the discussion of estimates from trade liberalization at the World Bank web site: http://www1.worldbank.org/wbiep/trade/TradePolicy.html.

TABLE 4
Testing the Natural Trading Partners Hypothesis

Equation	Welfare Change vs. (1)	Weighted Least Squares: Laspeyeres Index (2)	Weighted Least Squares: Fisher Ideal Index (3)
		Preferential Tariff Reduction	
(18)	Distance	−.16	−.08
		(.8)	(.15)
(19)	Distance and	−.16	−.05
		(.8)	(.15)
	Income	−1.80	−8.00
		(13.00)	(9.90)
(20)	Import volume	.50	.37
		(.65)	(.59)
		Preferential Tariff Reduction to Zero	
(18)	Distance	−.05	−.08
		(.13)	(.11)
(19)	Distance and	−.05	−.06
		(.14)	(.12)
	Income	.27	−6.58
		(10.03)	(7.99)
(20)	Import volume	−.06	−.29
		(.48)	(.50)

Note.—Figures in parentheses are standard errors. Distance is measured in thousands of miles, income in trillions of U.S. dollars, and volume of trade in billions of dollars. The left-hand side measuring welfare change in every case is 10^6 times the real income change (due to a unit reduction in tariff) measured per unit of U.S. income, i.e., $(Hdw/Y) \times 10^6$. Regressions were run using 1994 income and volume of trade data. For easy comparability, the second set of results, representing correlations with total preferential reduction in tariffs to zero, are per *unit* reduction in partner country tariffs.

preferential trade with "regional" partners), the following regression was run:

$$-\frac{HdW_j}{Y} = \alpha + \beta_1(\text{import volume})_j + \epsilon_j, \tag{20}$$

where import volume denotes bilateral import volume for the year 1994 (using bilateral import volumes from the beginning of the sample—1965—or from an intermediate year—1978—did not make any difference to the results).

Note that the dependent variable in each case was estimated and not observed. Further, the estimated standard errors associated with each of these observations on the dependent variable were different, raising the issue of heteroscedasticity. To correct for this, the method of weighted least squares (WLS) was used. The WLS estimates obtained by using the (inverse of the) estimated errors as weights are presented in table 4.

As these results indicate, the correlation between welfare change from preferential tariff reduction and distance is statistically insignificant. We cannot reject the null that distance does not matter. This can also be seen rather easily in figure 1, which plots the welfare estimates against distance (both adjusted by the heteroscedasticity correction described above). Clearly, no nonlinear relationship between these variables is revealed in these plots either.

As can be seen from table 4, the coefficient on distance remains insignificant even after conditioning on the partner's income level. Just as interesting, welfare changes appear to be uncorrelated with the volume of trade as well. Thus our tests are unable to find any evidentiary support for the natural trading partners idea in U.S. data. Figure 2, which plots welfare estimates against volume of trade (again correcting for the heteroscedasticity noted above), illustrates this. We should note here that if trade diversion was to be *ignored* and only trade creation to be considered in computing welfare effects, welfare changes would likely be *correlated* with the volume of trade. This should be clear from even a cursory examination of the estimates of own-price effects presented in tables 1 and 2. However, once the cross-price effects (which are negatively correlated with bilateral import volume) are taken into account, the correlation between overall welfare and trade volume disappears, as the results reported in table 4 reveal. Figure 3 plots the own-price effects and cross-price effects (for the Laspeyeres case) separately against trade volume, illustrating clearly the relevant positive and negative correlations just discussed.[11]

One issue arises because the estimates of welfare changes are, strictly speaking, valid for only small changes around the initial equilibrium. To see why this is important, assume that we start with equal tariffs against B and C. Assume that the own-price effect and the cross-price effect are estimated to be negative and positive, respectively. Now, note from (10) that even if the own-price effect was estimated to be larger in magnitude than the cross-price effect, the optimal tariff against the partner country (for any given positive tariff against the rest of the world) is positive (as can be seen by setting HdW equal to zero). It follows that lowering tariffs against the partner country below this optimal tariff implies a welfare loss relative to this optimum. (For example, tables 1 and 2 suggest that, when one starts from equal tariffs initially and holds tariffs against the rest of the world fixed, the optimal U.S. tariff against Canada is roughly a third of its tariff against the rest of the world; then reducing tariffs against Canada below this level would be welfare decreasing.) In principle, however, countries engaged in pref-

[11] These plots use raw estimates rather than values adjusted by the size of the corresponding standard errors since the correlations just discussed are a little easier to see in this case. However, WLS estimates of bivariate regressions between the own- and cross-price effects (separately) and trade volume do confirm these correlations strongly. The plots for the case with Fisher prices are very similar.

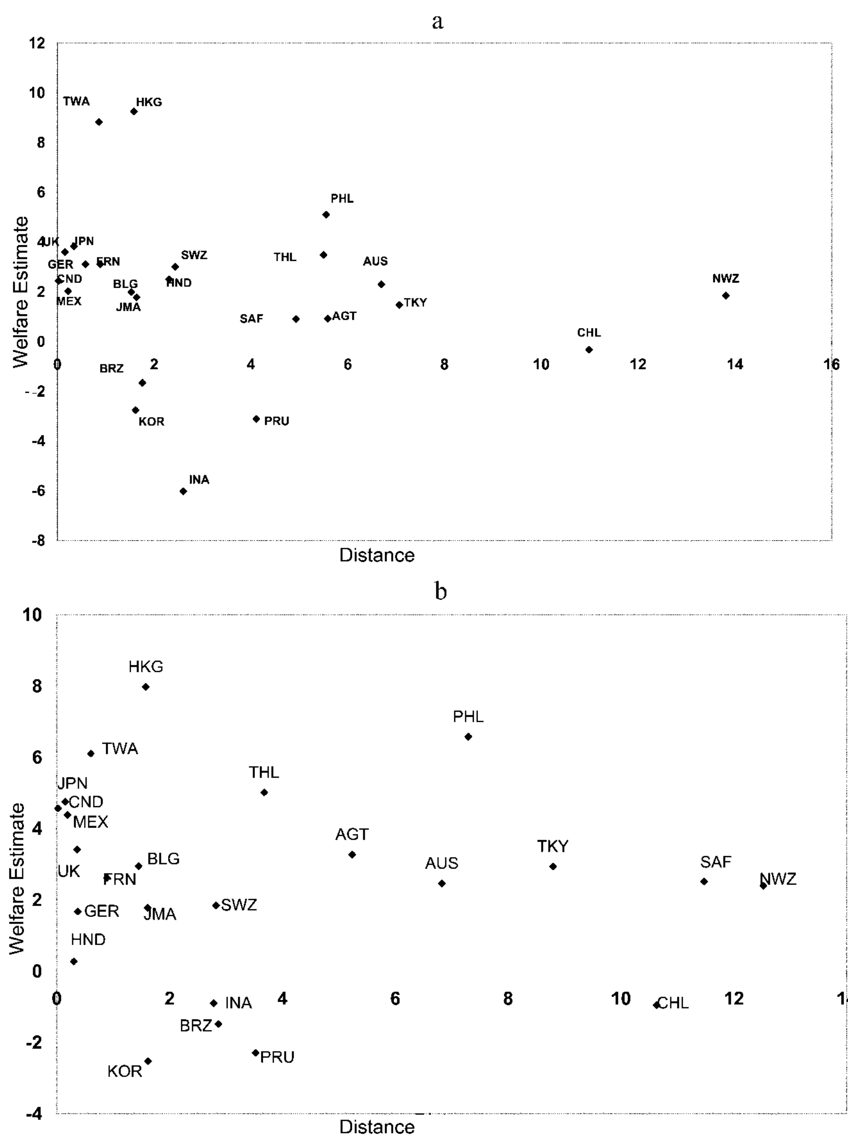

FIG. 1.—Welfare estimates vs. distance (heteroscedasticity adjusted). *a,* Laspeyeres. *b,* Fisher. Heteroscedasticity adjustment implies that the welfare estimates on the *y*-axis and the variables on the *x*-axis, distance and trade volume, have been divided by the corresponding standard error of the welfare estimates reported in tables 1 and 2. The units for all the original (i.e., unadjusted) variables are exactly as described in table 4.

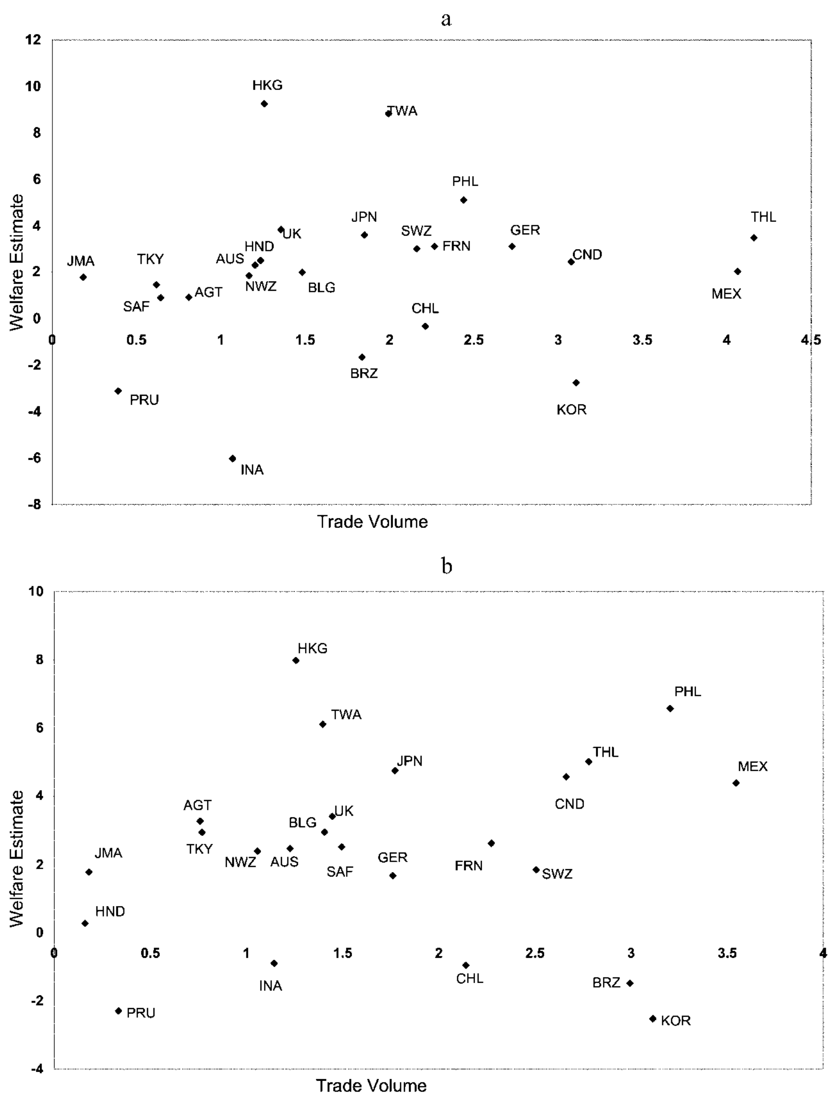

FIG. 2.—Welfare estimates vs. trade volume (heteroscedasticity adjusted). *a,* Laspeyeres. *b,* Fisher. Heteroscedasticity adjustment implies that the welfare estimates on the *y*-axis and the variables on the *x*-axis, distance and trade volume, have been divided by the corresponding standard error of the welfare estimates reported in tables 1 and 2. The units for all the original (i.e., unadjusted) variables are exactly as described in table 4.

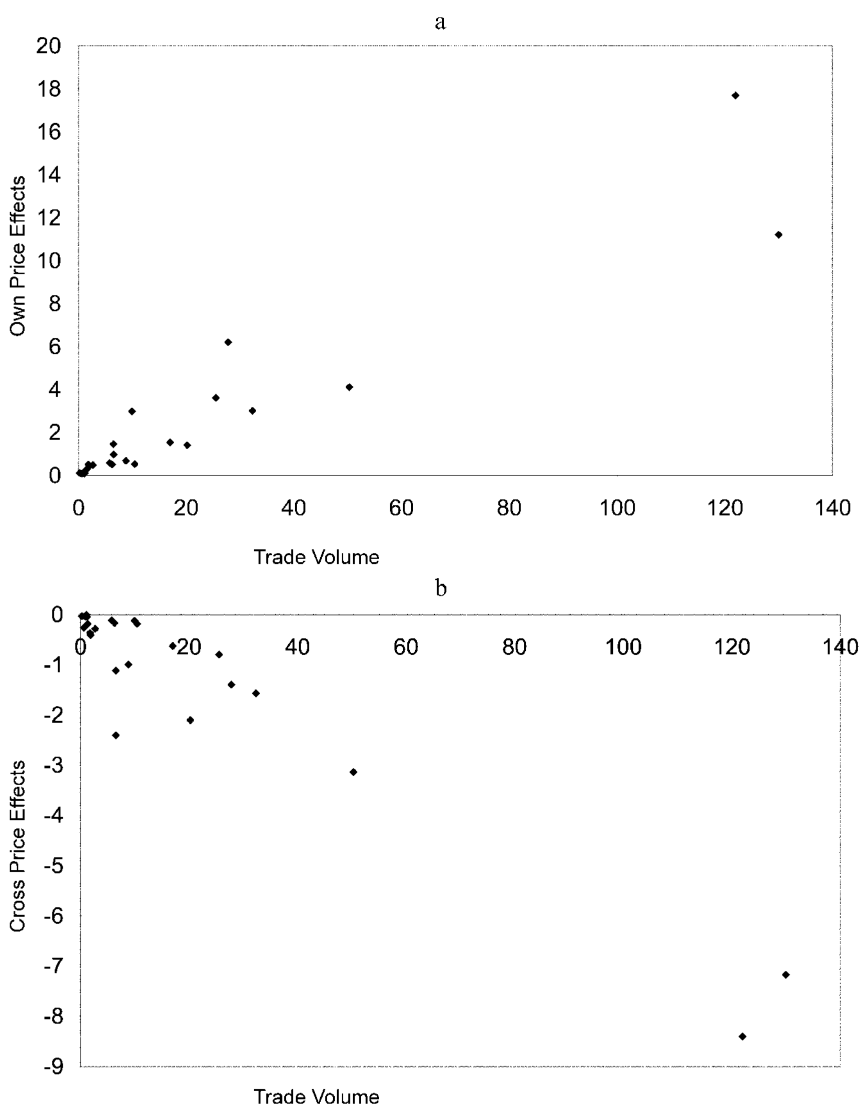

FIG. 3.—*a,* Own-price effects (Laspeyeres) vs. trade volume. *b,* Cross-price effects (Laspeyeres) vs. trade volume. Own- and cross-price effects are taken directly from table 1. Trade volume is measured in billions of U.S. dollars.

erential reduction in tariffs are required to reduce these preferential tariffs against the partner all the way to zero (as stipulated in Article XXIV of the GATT). Computing welfare changes from total preferential tariff reduction (to zero) would require us to take this into consideration.

Let us therefore consider whether our results would be significantly altered if we considered these "large" changes instead. It is useful to note, first, that the Slutsky own- and cross-price coefficients were estimated assuming that they are *invariant* to prices and income. Thus, to obtain welfare change due to a full reduction in partner country tariffs to zero, we can integrate both sides of (10) over the relevant interval. The calculation is considerably simplified by assuming that the term H and income are constant over this interval.[12] The estimates of "total" welfare change (again constructed taking starting tariffs against all countries as equal to the initial average U.S. tariff level in 1994) are reported in columns 2 and 3 of table 4. Regression results for (18), (19), and (20) with this *total* welfare change on the left-hand side instead are presented in table 4. The noncorrelation with distance remains, again, even after one controls for income levels. So does our finding of a noncorrelation with trade volume.

In summary, then, first, this paper obtains significant estimates of trade creation and trade diversion associated with preferential tariff liberalization against various partner countries. Second, the correlation between the overall welfare effect and distance is found to be statistically insignificant, and therefore the null that distance does not matter cannot be rejected—with and without conditioning on income levels. Nor can the null that the bilateral volume of trade does not matter. Our results are robust to the price measure used (the Laspeyeres or Fisher ideal index), time period chosen for the measure of partners' income level (in [19]), or the volume of trade (in [20]). In sum, our analysis does not find any evidentiary support for the natural trading partners theory in U.S. trade data.

VI. Conclusions

Many existing trade blocs are regional. Additionally, many possible extensions of existing trade blocs that are currently being discussed and

[12] To see that this is not a bad approximation for the purposes at hand, note that H is closely related to the inverse of the marginal utility of income and measures the actual expenditure incurred, at world prices, for a unit increase in welfare. Imports overall are about 10 percent of U.S. consumption. Of this, imports from any one country are less than about 20 percent (and only for the highest cases—Canada and Japan), implying that consumption of imports from any partner country is always less than about 2 percent of overall consumption.

debated in policy circles are forming, indeed, along regional lines, such as the expansion of NAFTA to include Chile, Argentina, and other South American countries or that of the European Union to include countries from eastern and central Europe. Academic and policy discussions on the future of the world trading system have focused substantially on the merits of such preferential trading arrangements.

This paper makes two contributions to this debate. Perhaps the primary contribution is a methodological one: The paper implements an estimation framework with tighter links to the underlying general equilibrium theoretical model and with far fewer restrictions as to functional forms than is traditional in the literature on preferential trade agreements. Second, in the actual implementation using U.S. trade data, this paper finds no support for the "natural trading partners" hypothesis—the idea that preferential tariff reduction with respect to geographically proximate countries is to be preferred over preferential treatment toward distant countries. Future research will tell whether this finding will be confirmed for other countries or in broader contexts even for the United States. The present analysis, nevertheless, points to the difficulty of identifying systematic economic criteria in choosing partners for preferential trading: Outcomes may simply be highly sensitive to context.

Several limitations of the economic analysis conducted here have been discussed in the text. Additionally, beyond trade and trade flows, it is worth mentioning other factors that have been suggested to be relevant in the decision to liberalize trade preferentially with one's neighbors—and that were ignored in the analysis. These include, inter alia, strategic defense considerations, the role of capital mobility, scale economies and the benefits of local economic concentration, the possibility of establishing stronger trading ties and networks with proximate countries due to, say, ethnic and cultural similarities, and so on. Serious consideration of these issues was entirely out of the scope of the present paper. They remain subjects for future research.

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