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An Empirical Model of Sunk Costs and the Decision to Export

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Firms that begin exporting face significant start-up costs. So, economywide exports respond differently to similar stimuli in different countries and time periods, depending on the number of firms that have already broken into foreign markets and the perceived permanence of export incentives.

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Summary findings

Exports respond unpredictably to a change in real exchange rates, suggests evidence from the 1980s.

Recent theoretical work (Paul Krugman and Richard Baldwin) explains this as a consequence of the sunk costs associated with breaking into foreign markets. Sunk costs include the cost of packaging, upgrading product quality, establishing marketing channels, and accumulating information on demand sources.

Roberts and Tybout use micro panel data to estimate a dynamic discrete-choice model of participation in export markets, a model derived from the Krugman-Baldwin sunk-cost hysteresis framework.

Applying the model to data on manufacturing plants in Colombia (1981–89), they test for the presence of sunk entry costs and quantify the importance of those costs in explaining export patterns. The econometric results reject the hypothesis that sunk costs are zero.

The results, which control for both observed and unobserved sources of plant heterogeneity, indicate that prior export market experience has a substantial effect on the probability of exporting, but its effect depreciates fairly quickly. The reentry costs of plants that have been out of the export market for a year are substantially lower than the costs of a first-time exporter. After a year out of the export market, however, the reentry costs are not significantly different from the entry costs.

Plant characteristics are also associated with export behavior: Large old plants owned by corporations are more likely to export than other plants.

Variations in plant-level cost and demand conditions have much less effect on the profitability of exporting than variations in macroeconomic conditions and sunk costs do. It appears especially difficult to break into foreign markets during periods of world recession.

This paper — a product of the International Trade Division, International Economics Department — is part of a larger effort in the department to describe the micro foundations of export supply response. The study was funded by the Bank's Research Support Budget under the research project "Micro-Foundations of Successful Export Promotion" (RPO 679-20). Copies of this paper are available free from the World Bank, 1818 H Street NW, Washington, DC 20433. Please contact Jennifer Ngaine, room R2-052, extension 37959 (37 pages). March 1995.

An Empirical Model of Sunk Costs and the Decision to Export

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Summary

The responsiveness of exports to changes in the incentive structure has long interested policy makers. But the empirical literature has provided little guidance on when or how exporters will respond to new incentive structures. Research in this area has produced a variety of supply elasticity estimates that vary dramatically across countries and time periods, and few hints on how to reconcile the diverse results.

Paul Krugman and Richard Baldwin have recently argued that the empirical literature fails because there are sunk costs associated with breaking into foreign markets -- including upgrading product quality, packaging, and the establishment of marketing channels. Hence the current-period export supply function depends upon the number and type of producers that were exporting in previous periods. Further, start-up costs mean that transitory policy changes or macro shocks can lead to permanent changes in market structure, and thus that trade flows may not be reversed when a stimulus is removed. That is, sunk entry or exit costs produce "hysteresis" in trade flows. Finally, when future market conditions are uncertain, sunk costs make patterns of entry and exit dependent upon the stochastic processes that govern variables like the exchange rate.

Taking the Baldwin/Krugman perspective as a point of departure, this paper develops an econometric model of a plant's decision to export. The model is fit to micro data for a large group of manufacturing plants in Colombia from 1981-1989, and used to directly examine the determinants of a plant's export decision for consistency with the theory.

The econometric results reject the hypothesis that sunk costs are zero. They also reveal that the re-entry costs of plants that have been out of the export market for a year are substantially less than the costs of a first-time exporter. Beyond a one year absence, however, the re-entry costs are not significantly different than those faced by a new exporter. This is consistent with the view that an important source of sunk entry costs for Colombian exporters is the need to accumulate information on demand sources, information that is likely to depreciate upon exit from the market.

While the results indicate that sunk costs are a significant source of export market persistence, both observed and unobserved plant characteristics also contribute to an individual plant's export behavior. For example, plants that are large, old, and owned by corporations are all more likely to export.

A number of policy implications emerge. For example, it appears especially difficult to break into foreign markets during periods of world recession. Also, although only 25 percent of the plants in the Colombian panel exported during the sample period, the results imply that sufficiently favorable macro conditions and/or reductions in sunk costs could make exporting profitable for a much larger proportion of plants. Finally, and most generally, the estimates imply that countries undertaking export promotion policies should distinguish measures aimed at expanding the export volume of exisitng exporters from policies aimed at promoting the entry of new exporters.

I. Introduction

Why is it that in some countries and time periods, a given trade and exchange rate regime supports large scale production for foreign markets, while in other countries or time periods, the same policies appear to induce a minimal export response? Put differently, why are estimates of export supply equations so sensitive to the time period or country under study?

In a recent series of papers Richard Baldwin, Paul Krugman, and Avinash Dixit have proposed an answer.[1] They begin from the assumption that non-exporters must incur a sunk entry cost in order to enter foreign markets. This makes the current-period export supply function dependent upon the number and type of producers that were exporting in previous periods. Further, it means that transitory policy changes or macro shocks can lead to permanent changes in market structure, and thus that trade flows may not be reversed when a stimulus is removed. That is, sunk entry or exit costs produce "hysteresis" in trade flows. Finally, when future market conditions are uncertain, sunk costs make patterns of entry and exit dependent upon the stochastic processes that govern variables such as the exchange rate. Under plausible assumptions, greater uncertainty makes trade flows less responsive to changes in these variables. None of these implications of sunk costs is captured in standard empirical export supply functions, and all could contribute to the "instability" of empirical relationships.[2]

To date, attempts to empirically validate the sunk-cost hysteresis framework have focussed on

[1] See, in particular, Dixit, 1989a and 1989b; Baldwin, 1988 and 1989; Baldwin and Krugman, 1989; Krugman, 1989).

[2] In their review of empirical studies of price and income elasticities for traded goods, Goldstein and Khan (1985, pp. 1087-1092) report a very wide range of estimates for the supply elasticity of total exports from developed countries. They conclude that "excluding the United States, the supply-price elasticity for the total exports of a representative industrial country appears to be in the range of one to four. The supply elasticity for U.S. exports is probably considerably higher than that, perhaps even reaching ten to twelve." They also discuss some evidence indicating that the response of export supply to price changes is slower than demand-side adjustments. They speculate that this may reflect start-up costs associated with export production or greater uncertainty associated with selling abroad.

asymmetries in the response of trade flows to exchange rate appreciation versus depreciation.[3] A limitation of this approach is that data on the volume of trade flows, even for very disaggregated commodities, cannot distinguish the entry and exit of exporters from the supply response of continuing exporters. With the exception of Campa (1993), the foreign market entry and exit patterns, which are the focus of the theory, have not been examined for consistency with the sunk-cost hysteresis model.[4]

In this paper we develop an empirical test of the sunk-cost hysteresis model that directly examines entry and exit patterns in plant-level panel data. We develop and estimate a dynamic discrete choice model of the decision to export when sunk entry or exit costs are present. In essence, this model predicts exporting status in the current period as a function of plant characteristics, previous exporting status and a serially correlated disturbance. It not only permits us to formally test for the presence of sunk costs (using coefficients on lagged exporting status), it allows us to summarize the effects of time, individual producer characteristics, and prior exporting experience on the probability of participating in the export market. The data we use describe the export patterns of

[3] The empirical evidence derived from trade flow data has produced no clear consensus. Based on aggregate U.S. data, Baldwin (1988) concludes that the substantial appreciation of the U.S. dollar during the early 1980's resulted in a structural shift in U.S. import pricing equations. This is consistent with sunk-cost hysteresis. In contrast, Gagnon (1987) finds that trade has been more responsive to relative prices in the more uncertain post-Bretton Woods era, a result inconsistent with some versions of the hysteresis model. Using time-series data for U.S manufacturing industries, Feinberg (1992) finds that exports became more dispersed across destination markets as the dollar depreciated, suggesting that there was firm entry into new country markets. The effect was weaker in industries where distribution networks, and thus presumably sunk entry costs, are more important. Parsley and Wei (1993) focus on bilateral U.S.-Canada and U.S.-Japan trade flows for very disaggregated commodities. They find that both the past history of U.S. exchange rate changes and measures of exchange-rate volatility had no significant effect on trade flows. Both findings are inconsistent with the hysteresis model.

[4] Campa (1993) examines the number of foreign firms that made direct investments in the 61 U.S. wholesale trade industries over the 1981-87 period. He finds that exchange-rate uncertainty, which is proxied by the standard deviation of the monthly rate of growth of the exchange rate, is negatively correlated with the number of firms investing in the U.S.. He also reports that an industry's sunk costs, which are proxied by the advertising-sales ratio and ratio of fixed assets to net worth of firms in the industry, is negatively correlated with foreign-firm entry. Both findings are consistent with the hysteresis model. Although not in a trade context, related work by Bresnahan and Reiss (1991, 1994) shows how data on net entry into a market can be used to make inferences about the ratio of sunk entry and exit costs to average profitability. Their technique exploits the asymmetric response of the number of producers to population (demand) changes across different geographic markets. However, as they acknowledge, persistance in behavior due to permanent cross-producer differences in profitability can create the appearance of sunk costs in their model.

Colombian manufacturing plants in four major exporting industries over the period 1981-1989, a nine-year span characterized by substantial changes in aggregate demand and real exchange rates.

The empirical results strongly reject the hypothesis that sunk costs are zero. This implies that prior export market experience significantly affects the current decision to export. Further, although recent experience in foreign markets is extremely important, its effect depreciates fairly quickly over time. A plant that exported in the prior year is up to 40 percentage points more likely to export in the current year than an otherwise comparable plant that has never exported. But by the time a plant has been out of the export market for two years its probability of exporting differs little from that of a plant that has never exported.

Several policy implications emerge. First, although only 25 percent of the plants in our panel exported during the sample period, our results imply that sufficiently favorable macro conditions and/or reductions in sunk costs could make exporting profitable for a much larger proportion of plants. Second, our estimates imply that countries undertaking export promotion policies should distinguish measures aimed at expanding the export volume of exisitng exporters from policies aimed at promoting the entry of new exporters.

In the next section of the paper we summarize the theoretical sunk-cost model. The third section provides an overview of the patterns of export participation among Colombian manufacturing plants between 1981 and 1989. The fourth section develops an econometric model of the export decision, and the fifth section presents our results. We briefly summarize and draw conclusions in the sixth section. Readers uninterested in methodological issues may wish to skip section II and readers uninterested in econometric problems may wish to skim section IV.

II. A Theoretical Model of Entry and Exit with Sunk Costs

As reviewed in Krugman (1989), sunk costs affect the export supply function for several reasons. First, and most obviously, once the sunk costs of entering a market have been met, a producer will remain in that market as long as operating costs are covered. This implies that changes in policy, exchange rates, or prices in foreign markets can permanently alter market structure and thus observed export behavior. For example, devaluations that induce entry into the export market may permanently increase the flow of exports, even if the currency subsequently appreciates. Second, even if current conditions appear favorable to exporting, they may not induce entry into the export market if they are regarded as transitory. In this case, the expected future stream of operating profits may not cover the sunk costs of entering foreign markets. Thus large devaluations may induce little response from potential exporters if they are perceived as transitory. Finally, as formally demonstrated by Dixit (1989a), the combination of sunk costs and uncertainty about future market conditions can create an option value to waiting. Dixit's simulation results suggest that even small amounts of uncertainty can significantly magnify the degree of persistence in a producer's exporting status.

To motivate our empirical work, we begin by reviewing the theoretical models that generate these results (see footnote 1 for references). For each period t, let the i^{th} plant's expected gross profits when exporting differ from its expected gross profits when not exporting by the amount $\pi_i(p_t,s_{it})$. Here p_t is a vector of market-level forcing variables that the plant takes as exogenous (e.g., the exchange rate), and s_{it} is a vector of state variables specific to the plant (e.g., capital stocks and geographic location). Once in the market, plants are assumed to freely adjust export levels in response to current market conditions (Baldwin, 1989). Thus the function $\pi_i(p_t,s_{it})$ represents the increment to expected profits associated with exporting in year t, assuming that the profit-maximizing level of exports is always chosen.

These profits are gross because they have not been adjusted for the sunk costs of foreign market entry or exit. Assume that if the i^{th} plant last exported in year t-j $(j \geq 2)$ it faces a re-entry cost of F_i^j, so upon resuming exports in year t it earns $\pi_i(p_t, s_{it}) - F_i^j$. Similarly, if the plant had never exported previously, it faces an entry cost of F_i^0 and earns $\pi_i(p_t, s_{it}) - F_i^0$ in its first year exporting. Finally, a plant that exported in period *t-1* earns $\pi_i(p_t, s_{it})$ during period *t* by continuing to export and $-X_i$ if it exits. As in Dixit (1989a), these sunk costs represent the direct monetary costs of entry and exit. The *j* superscript generalizes previous models to allow sunk re-entry costs to depend on the length of absence from the market. This could reflect the increasing irrelevance of the knowledge and experience gained in earlier years, or the increasing cost of updating old export products. The *i* subscript allows sunk costs to vary across plants with differences in size, location, previous experience, and other plant characteristics.[5]

To collapse these earnings possibilities into a single expression, define the indicator variable Y_{it} to take a value of 1 if the plant is exporting in period *t*, and 0 otherwise. Also, let the exporting history of the plant through period *t* be given by $Y_{it}^{(-)} = \{Y_{it-j} \mid j=0 \ldots J_i\}$, where J_i is the age of the plant. Then period *t* exporting profits are:

$$R_{it}(Y_{it}^{(-)}) = Y_{it}\left[\pi_{it} - F_i^{0}(1-Y_{it-1}) - \sum_{j=2}^{J_i}(F_i^j - F_i^{0})\tilde{Y}_{it-j}\right] - X_i Y_{it-1}(1-Y_{it}) \ ,$$

where $\tilde{Y}_{it-j} = \left[Y_{it-j}\prod_{k=1}^{j-1}(1-Y_{it-k})\right]$. This last expression summarizes the plant's most recent exporting experience: $\tilde{Y}_{it-j} = 1$ when the plant's most recent exporting experience occurred *j* years earlier and 0 otherwise.

In period t, managers are assumed to choose the infinite sequence of values $Y_{it}^{(+)} = \{Y_{it+j} \mid j \geq 0\}$ that maximizes the expected present value of payoffs. In period *t*, this maximized payoff is:

[5] To keep the notation tractable we have not added a time subscript to entry and exit costs. In the empirical section, we will test whether they vary over time, as would be expected if there are changes in credit market conditions or trade policies that affect access to foreign markets.

$$V_{it}(\Omega_{it}) = \max_{Y_{it}^{(\cdot)}} E_t \left[\sum_{j=t}^{\infty} \delta^{j-t} R_{it} \mid \Omega_{it} \right]$$

where δ is the one-period discount rate and expectations are conditioned on the plant-specific information set, Ω_{it}. Using Bellman's equation, plant i's current exporting status can be represented as the Y_{it} value that satisfies:

$$V_{it}(\Omega_{it}) = \max_{Y_{it}} \left(R_{it}(Y_{it}^{(-)}) + \delta E_t \{ V_{it+1}(\Omega_{it+1}) \mid Y_{it}^{(-)} \} \right)$$

where E_t denotes expected values conditioned on the information set Ω_{it}. From the right-hand side of this expression, it follows that the i^{th} plant will be in the export market during period t if:

$$\text{(1)} \quad \begin{aligned} &\pi_i(p_t, s_{it}) + \delta\left[E_t\left(V_{it+1}(\Omega_{it+1}) \mid Y_{it}=1\right) - E_t\left(V_{it+1}(\Omega_{it+1}) \mid Y_{it}=0\right)\right] \geq \\ &\qquad F_i^{\circ} - (F_i^{\circ} + X_i) Y_{it-1} + \sum_{j=2}^{J_i} (F_i^{\circ} - F_i^j) \tilde{Y}_{it-j} \end{aligned}$$

where $-(F_i^{\circ} + X_i)$ is the sum of sunk entry costs for a plant that never exported and exit costs for current exporters, sometimes referred to as the "hysteresis band" (Dixit, 1989a).

Equation (1) provides the participation condition that will be estimated in section V. It has several empirical implications we will pursue. First, if there are no sunk costs, the participation condition collapses to $\pi_i(p_t, s_{it}) \geq 0$. Hence one can test the sunk-cost hysteresis framework by asking whether, given a plant's current gross profits, its exporting history helps explain its current exporting status. Second, if sunk costs *do* matter, equation (1) implies that they appear directly in each plant's participation condition as coefficients on binary variables that describe its exporting history.[6] Hence the magnitude of sunk costs and the rate at which past experience depreciates can be identified.

[6] The influence of sunk costs also comes through the expected value term on the left-hand side, but since this is a non-linear expression, coefficients on the indicator variables are identified.

Finally, this equation indicates that realizations on the variables p_t and s_{it} influence export decisions through their effect on $\pi_t(p_t, s_{it})$ *and* their effect on the expected future value of becoming an exporter now. This latter effect implies, for example, that exchange rate movements that managers consider transitory will generally have less effect than equivalent movements that are viewed as long-term regime shifts.

III. The Pattern of Export Participation in Colombia

Before discussing estimation issues and results, it is useful to introduce our data base, review the export environment in Colombia, and provide some aggregate evidence on the pattern of export market participation during the 1980s.

The Data: The analysis in this paper is based on annual plant-level data collected as part of the Colombian manufacturing census for the years 1981-1989. This census, which covers all plants with 10 or more employees, provides information on each plant's geographic location, industry, age, ownership structure, capital stocks, investment flows, expenditure on labor and materials, value of output sold in the domestic market, and value of output exported. We have matched the individual plant observations across years to form a panel.[7] The data are particularly well-suited to analyzing export market participation because they allow us to observe transitions of individual plants into and out of the export market and to control for some important observable plant characteristics that are likely to affect the export decision.

The Policy Regime and Export Participation Rates: Table 1 illustrates the basic patterns

[7] The census data and matching process are discussed in greater detail in Roberts (1994).

over the sample period for the 19 major exporting industries in the Colombian manufacturing sector.[8] In general terms, the macro environment in Colombia was not conducive to profitable exporting of manufactured goods in the early 1980's. Responding to illegal exports, foreign capital inflows, and a boom in the coffee market, the Colombian peso appreciated steadily between the mid-1970's and 1983. As shown in Table 1, this pattern was reversed after 1983, with the currency losing approximately one-half of its value by 1989. This partly reflected central bank currency market interventions to ease competitive pressures on tradeable goods producers.

The time-series pattern of manufactured exports largely mirrors this movement in the exchange rate. The real value of manufactured exports from the nineteen major exporting industries declined slightly from 109.6 (billion 1985 pesos) in 1981 to 95.9 in 1984, for an average annual growth rate of -4.45 percent. From 1984 to 1989 the pattern reversed and the quantity of exports grew at an annual average rate of 21.1 percent.[9]

Commercial policy sheltered import-competing producers throughout the sample period. Substantial tariff barriers were reduced slightly after 1984, but quantitative restrictions on the imports of products that competed with domestic industries were maintained. In addition to turning the terms of trade against exporters, these polices made it more difficult to import raw materials or capital goods that may have been necessary to increase the quality of manufactured products.[10]

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[8] The nineteen industries and their SITC codes are: food processing (311/312), textiles (321), clothing (322), leather products (323/324), paper (341), printing (342), chemicals (351/352), plastic (356), glass (362), non-metal products (369), iron and steel (371), metal products (381), machinery (382/383), transportation equipment (384), and miscellaneous manufacturing (390). These industries account for over 96 percent of Colombia's manufacturing sector exports and 85 percent of manufacturing output in each sample year.

[9] The real value of exports is measured as the peso value of exports deflated by an export price index from the IMF International Financial Statistics. This measure will overstate the dependence of the physical volume of exports on the exchange rate because of valuation effects.

[10] The long-term protection of the domestic market from import competition appears to have also contributed to the low product quality and low productivity that have made it difficult for Colombian exporters to compete in the international market (World Bank, 1992).

Nonetheless, the bias toward import-competing activities was partly offset by export subsidies, which increased relative to the value of exports by approximately 50 percent between 1983 and 1984, and thereafter declined.[11] The incentives to export created by export policy therefore were counter to those created by exchange rate movements.

The net effect of these changes on the number of exporting plants and the proportion of plants that exported is summarized in the last two rows of Table 1. Again, the time-series pattern largely reflects the movement in the exchange rate. Through 1984, there was net exit from the export market, and a decline in the proportion of plants exporting. After that year there was net entry and a steady increase in the participation rate. There is, however, some evidence of asymmetry in the magnitude of the response. The modest 6 percent real appreciation between 1981 and 1983 was accompanied by a decline in the export participation rate from .129 to .113, but a much larger (45 percent) depreciation between 1984 and 1989 served only to increase the participation rate to .135. From this short time series it appears that it took both a substantial and persistent devaluation to induce entry into the export market. This is consistent with the conjecture that potential exporters faced substantial start-up costs.

Survey evidence further supports this hypothesis.[12] First, to sell in developed country markets, Colombian producers were often required to invest in product quality upgrading. Second, there was little exporting infrastucture in the form of trading companies or distribution agents. These companies typically provide transportation, customs, and shipping services, as well as information on

[11] This reduction in export subsidies reflected a reduction in two government programs used to promote exports. The first program rebates customs duties paid on imported materials and capital equipment for plants that export. In 1980 41 percent of the value of exports came from plants that received rebates. This rose to 62 percent in 1984 and fell to 53 percent in 1986. The second program provides direct subsidies to exporters based on the value of their exports. The average rate of subsidy increased from 7.6 percent of the value of exports in 1981 to 15.0 percent in 1985 and then fell to 8.7 percent in 1986.

[12] The discussion in this section is based on World Bank (1992), which summarizes and interprets interviews with the managers of several hundred Colombian plants.

prices, potential buyers, and product standards or requirements in other countries. The absence of these middlemen probably discouraged potential exporters, both by increasing the information costs they faced, and by increasing the degree of uncertainty concerning foreign market conditions. Apparently, however, the lack of a well-developed trading services sector did not affect all producers equally. Exporters able to deal in large volumes or to ship to large markets were relatively less constrained by the absence of trading intermediaries because they were able to sell directly to final buyers. This finding suggests that sunk costs rose less than proportionately with export volume.

Export market entry was also inhibited by institutional factors that affected expected profits. Notably, a survey of Colombian financial institutions revealed that none were willing to lend money against export orders or letters of credit from purchasers' banks. Lenders attributed this unusually conservative practice to their inability to judge whether the potential borrowers could seriously compete abroad. It mainly hurt first-time exporters and existing one-product, one-country exporters attempting to enter new country or product markets.

Finally, as emphasized by Dixit (1989a), regime uncertainty may have induced producers to delay entry into the export market, even after substantial devaluation. In the World Bank survey, producers cited uncertainty about the permanence of the change in trade and exchange rate regimes as incentives to delay or forego entry into the export market. Their main concern was apparently that lobbyists in favor of protecting domestic industries would be able to reverse the trend toward trade liberalization.

In summary, during the sample period many Colombian manufacturers viewed the export market as more risky and less profitable than the domestic market. The lack of a trading services sector, access to financing, and low product quality all appear to have constrained export market participation by raising the costs of entry or increasing the uncertainty of the profitability of exporting.

Entry and Exit in the Export Market: The analytical model reviewed in Section II implies that this combination of sunk costs and uncertainty should induce persistence in producers' exporting status. That is, those who have already incurred the sunk start-up costs should be relatively likely to export in the current period. Some preliminary evidence on this prediction is provided by transition rates into and out of the export market, which are summarized in Table 2. Each row describes a transition from the exporting status in column 1 to the status in column 2. The entries in the table are the proportion of plants in each of the period t categories that choose each of the two possible categories in year $t+1$.[13] The top panel applies to the 19 major manufacturing industries, and the bottom panel applies to the 650 plants in the four major exporting industries -- food, textiles, paper, and chemicals -- that will be used to estimate the econometric model in the next section.

The top row of each panel indicates that, of the plants that did not export in year t, more than 95 percent of them did not export in year $t+1$. For the plants initially in the export market, the proportion of manufacturing plants that remain in the market from one year to the next varies from 83 percent to 91 percent over time, and the proportion of plants in our four-industry subsample that remains in the market varies from 85 percent to 95 percent. Clearly, there is substantial persistence in the plant-level patterns of export market participation. Nonetheless, only 36 percent of the plants in the subsample that exported at some time remained exporters for the whole sample period, and among plants that *did* change exporting status, 60 percent did so more than once.[14]

Persistence in exporting status might be caused by sunk costs, as the hysteresis models suggest. Alternatively, it might be caused by underlying plant heterogeneity: persistent cross-plant

[13] The data correspond to the group of plants that were in operation in each year 1981-1989. There are 2369 plants in this group and they represent approximately 40 percent of the number of plants in operation in any year. On average, 18.1 percent of these plants participated in the export market in any single year and they accounted for 62.3 percent of the total number of exporters and 61.6 percent of the value of manufactured exports.

[14] The appendix summarizes the patterns of multiple switches into and out of the export market and the ability of the empirical model to explain these patterns.

differences in the payoff from exporting, $\pi_i(\cdot)$, would explain why some plants are always in the export market and others are always out. Similarly, the fact that many exporters enter or exit the market multiple times can also be interpreted several ways: it could mean that sunk costs are small, or it could reflect lingering benefits from having exported recently. In the next section we develop an econometric framework that can discriminate among these competing explanations.

IV. An Empirical Model of Export Market Participation

The Estimating Equation: Our empirical model of a plant's exporting decision begins with the participation condition given by equation (1). Define

$$\pi_{it}^{*} = \pi_i(p_t, s_{it}) + \delta\left[E_t\left(V_{it}(\Omega_{it+1}) \mid Y_{it}=1\right) - E_t\left(V_{it}(\Omega_{it+1}) \mid Y_{it}=0\right)\right]$$

as the latent variable representing the expected increment to gross future profits for plant i if it exports in period t. Export market participation is then summarized by the dynamic discrete choice equation:[15]

$$(2) \qquad Y_{it} = \begin{cases} 1 \text{ if } & \pi_{it}^{*} - F_i^{o} + (F_i^{o} + X_i)Y_{it-1} + \sum_{j=2}^{J_i} (F_i^{o} - F_i^{j})\tilde{Y}_{it-j} \geq 0 \\ 0 \text{ otherwise} & \end{cases}$$

There are two ways we might proceed to estimate equation (2). First, we could develop a structural representation of the participation condition by making specific assumptions about the form of the profit function and the processes that generate s_{it} and p_t.[16] Alternatively, we could forego identification of structural parameters, and approximate $\pi_{it}^{*} - F_i^{o}$ as a reduced-form expression in exogenous plant and market characteristics that are observable to producers in period t. The advantage of the first approach is that, in principle, it allows identification of the parameters of the profit function (*inter alia*) and provides a complete description of the dynamic process. Its main disadvantage is that very restrictive parameterizations are required to make structural estimation feasible. This problem is particularly acute in our model because the dependence of sunk entry costs upon the length of time out of the export market implies a participation series that is a J^{th}-order

[15] This equation is similar to those used to study labor market participation and employment (Heckman, 1981a, 1981b).

[16] Eckstein and Wolpin (1989) and Rust (1993) summarize the literature on estimating structural dynamic models of discrete choice.

Markov process. Because of this difficulty, and because we do not need a structural model to assess the role of sunk costs or to investigate the sensitivity of decisions to s_{it} and p_t, we pursue the reduced-form approach.

To parameterize the reduced-form model, we assume that variation in $\pi^*_{it} - F^o_i$ arises from three different sources: time-specific effects that reflect industry or macro-level changes in export conditions (μ_t), observable differences in plant characteristics (Z_{it}), and noise (ϵ_{it}):

$$\pi^*_{it} - F^o_i = \mu_t + \beta Z_{it} + \epsilon_{it} \tag{3}$$

The term μ_t is an annual time effect reflecting temporal variations in export profitability and start-up costs that are common to all plants. These time effects pick up the influence of credit market conditions, exchange rates, trade policy conditions, and other time-varying factors captured by p_t in the analytical model. The vector Z_{it} controls for factors represented by s_{it} and F^o_i in the analytical model: exogenous plant-specific determinants of current operating profits and start-up costs. It includes a constant, a set of industry dummies defined at the three-digit SITC level, a dummy variable to control for the ownership structure of the plant (proprietorship and partnership versus corporation), and a set of two locational dummies to distingish the Bogota and Medillin/Cali regions from all others.[17] The vector Z_{it} also includes several continuous variables lagged one period and measured in logarithms: the ratio of foreign to domestic prices for output, the wage rate, capital stock, and plant age.[18] Relative prices and wages affect the attractiveness of domestic versus foreign markets. Capital stock and age proxy for efficiency: in addition to scale effects, studies of industrial evolution

[17] The base group for comparison is a plant in the food industry that is not a corporation and that is located in the Bogota area.

[18] Foreign prices are constructed using unit values of exports at the four-digit ISIC level; domestic prices were obtained at the same level from the Central Bank of Colombia.

suggest that efficient producers are more likely to survive and grow.

Additional restrictions on sunk entry and exit costs are needed to identify the model. Let $\gamma_i^j = F_i^o - F_i^j$ $(j = 2, \ldots, J)$ and $\gamma_i^j = \gamma_i^o = F_i^o + X_i$ $(j \geq J+1)$, implying that experience is completely depreciated if it was acquired more than J years ago. (The problem of choosing J will be discussed later). Further, let $\gamma_i^j = \gamma^j$, implying that cross-plant variation in sunk costs is negligible among producers who acquired their most recent exporting experience at the same time. Then substituting (3) into (2), we obtain our basic estimating equation:

$$
(4) \qquad Y_{it} = \begin{cases} 1 \text{ if } 0 \leq \mu_t + \beta Z_{it} + \gamma^o Y_{it-1} + \sum_{j=2}^{J} \gamma^j \tilde{Y}_{it-j} + \epsilon_{it} \\ 0 \qquad \text{otherwise,} \end{cases}
$$

Properties of the disturbance term ϵ_{it} will be discussed in the following subsection.

As noted earlier, the participation decision does not depend upon exporting history if sunk costs are zero. Hence we can test the null hypothesis that sunk costs are unimportant in the export decision by testing whether γ^o and γ^j's are jointly equal to zero. If they are significant, we can use them to make inferences about the rate at which export market experience decays. Using interaction terms between the lagged participation variables and plant characteristics or macro variables, the model can also be generalized to allow the sunk cost parameters γ's to vary with changes in these variables. Finally, we can use equation (4) to study the importance of temporal (μ_t) and cross-plant (βZ_{it}) variation in net expected profits from exporting ($\pi_{it}^* - F_i^o$). In particular, we can impute probabilities of entry or exit in response to a given shift in exogenous variables for plants with different characteristics.

Econometric Issues: To isolate the importance of sunk costs, it is critical that we control for all other sources of persistence in exporting status. Much of this task is accomplished by including the vector of observable plant characteristics Z_{it} in equation (4). However, it is very likely that some characteristics, such as managerial expertise or output quality, will remain unobserved and their presence will induce serial correlation in the error term, ϵ_{it}. If we use an estimator that ignores this serial correlation, the model will incorrectly attribute its effect on exporting status to past participation and thus overstate the importance of sunk costs.[19]

Following Heckman (1981a, 1981b) we allow for serial correlation by assuming that ϵ_{it} is the sum of a permanent, plant-specific component and a white-noise component: $\epsilon_{it} = \alpha_i + \omega_{it}$. Here α_i represents unobservable plant-level differences in managerial efficiency, foreign contacts, and other factors that induce persistent plant-specific differences in the returns from exporting. We normalize $var(\epsilon_{it}) = 1$, and assume that $cov(\alpha_i,\alpha_k) = 0 \;\; \forall\, i \neq k$, $cov(Z_{it},\epsilon_{it}) = cov(\alpha_i,\omega_{it}) = 0 \;\; \forall\, i, t$, and $cov(\omega_{it},\omega_{i,t-j}) = 0 \;\; \forall\, j \neq 0$. This specification implies that $cov(\epsilon_{it},\epsilon_{it-j}) = var(\alpha_i) = \sigma^2_\alpha$, so the parameter σ^2_α is both the covariance between different time periods for a single plant, and the fraction of the variance of ϵ_{it} that arises from the permanent component in the error. Assuming that α and ω are each normally-distributed random variables, equation (4) can be estimated as a random-effects probit model.[20]

There remains an additional problem. We observe a plant's export status in years 1 through T, and our lag structure reaches back J periods, so equation (4) can be used to model the export

[19] This is the problem of "spurious state dependence" discussed in the empirical literature on labor market participation. See, for example, Heckman (1981a).

[20] We do not control for the α_i by using plant-specific dummy variables because of the "incidental parameters problem" discussed in Neyman and Scott (1948), Chamberlin (1980), and Heckman (1981c). For a given number of time periods, the number of α_i values grows in direct proportion to the sample size, making consistent estimation as $n \rightarrow \infty$ more difficult. Under these conditions, a standard logit or probit estimator using plant-specific dummy variables will not yield consistent slope coefficients. If the time dimension of the panel is small and/or the model is dynamic, the bias can be substantial. See Hsiao(1986, pp. 159-161), Wright and Douglas (1975) and Heckman (1981c), for discussion of the magnitude of the bias. In particular, Heckman (1981c) finds that the bias in slope coefficients from a *dynamic* probit with unobservable effects is "disturbingly large" (p. 180) when T = 8.

decision in years $J+1$ through T. But Y_{it-1} and $\tilde{Y}_{it-j}$ values corresponding to these first J years cannot be treated as exogenous determinants of Y_{it} $(J+1 < t \leq 2J)$ because each depends on α_i. Heckman (1981c) suggests dealing with this "initial conditions" problem by using an approximate representation for Y_{it} when $t \leq J$ and allowing the disturbances in the first J periods to be correlated with the disturbances in every other period. Specifically, suppose that expected profits in the export market during periods 1 through J can be represented with the equation:

$$(5) \qquad \pi^*_{it} - F^o_i = \lambda \bar{Z}_{it} + \varphi_{it} \qquad\qquad t = 1, \ldots J$$

where $\bar{Z}_{it}$ is a vector of pre-sample information from periods prior to t on the i^{th} plant, λ is a conformable parameter vector and the random variable φ_{it} is assumed to be normally distributed with $var(\varphi_{it}) = 1$, $cov(\varphi_{it}, \omega_{it}) = 0$.[21] Critically, we also let φ_{it} be correlated with the persistent plant component of the error term ϵ_{it}: $cov(\varphi_{it}, \alpha_i) = \rho$. (Note that ρ can also be interpreted as the fraction of variation due to α_i in the first J periods.) Then the correlation of lagged dependent variables with α_i in years 1 through J can be recognized in the likelihood function by representing the Y_{it} process using equation (5) for periods $t=1,\ldots,J$, and equation (4) for periods $t=J+1,\ldots,T$. This specification adds the nuisance parameter vector λ to the model as well as the parameter ρ. Although equation (5) is an imperfect representation of the process generating the data, simulation evidence suggests that Heckman's procedure performs reasonably well.[22]

To construct the likelihood function for observations in periods 1 through T define:

[21] In the empirical work we include all of the plant characteristics in Z_{it} described above as explanatory variables in the initial-conditions equation (5). Also included are two-year lagged values of the plant's wages, capital stock, and export price.

[22] An alternative solution to this problem is prevented because of the presence of time-varying exogenous variables in the model. These make it impossible to solve the model for the steady-state probabilities of the pre-sample Y_{it} realizations as functions of data and estimable parameters.

(6) $b_{it} = -[\,\mu_t + \beta Z_{it} + \gamma^{\circ} Y_{it-1} + \sum_{j=2}^{J} \gamma^j \tilde{Y}_{it-j} + \xi_1 \alpha_i],$ $\qquad t = J+1, \ldots T$

and

(7) $a_{it} = -(\lambda \bar{Z}_{it} + \xi_2 \alpha_i),$ $\qquad t = 1, \ldots J$

where $\xi_1 = (\sigma_\alpha^2 / (1 - \sigma_\alpha^2))^{1/2}$, and $\xi_2 = (\rho/(1-\rho))^{1/2}$. The likelihood function for a panel of n plants can now be written as:

(8) $$L(\lambda, \mu_1, \ldots \mu_T, \beta, \gamma^{\circ}, \gamma^2, \ldots, \gamma^J, \sigma_\alpha^2, \rho)$$

$$= \prod_{i=1}^{n} \int_{-\infty}^{\infty} \{ \prod_{t=1}^{J} \Phi[a_{it}\,(2Y_{it}-1)] \prod_{t=J+1}^{T} \Phi[b_{it}\,(2Y_{it}-1)]\} \,\phi(\alpha)\, d\alpha.$$

Here $\Phi(\cdot)$ is the standard normal cumulative distribution function and $\phi(\alpha)$ is the normal density function for the unobserved plant effects.

To summarize, the model of export market participation consists of equations (4) and (5) and is estimated with maximum likelihood using the likelihood function in equation (8).[23] The specification of the error terms allows for serial correlation arising from plant-specific differences in the profitability of exporting and for the initial conditions problem. Both cross-sectional and temporal variation in the data are used to identify the coefficients. The former is due mainly to cross-plant differences in industry, location, business type, age, capital stock, the relative price of foreign to domestic output, and wages. The latter is due largely to economy-wide fluctuations in macro conditions, and to unobserved plant-specific shocks (φ_{it} and ω_{it}), which combine with changes in Z_{it} to induce temporal variation in Y_{it}. Note that, even if there were no variation in Z_{it}, the transitory

[23] The integral in equation (8) is approximated using the Hermite integration formula discussed by Butler and Moffit (1982). The results reported below used seven evaluation points and were virtually unchanged from a five-point evaluation.

shocks φ_{it} and ω_{it} would suffice to identify the coefficients on Y_{it-1} and $\tilde{Y}_{it-j}$.

Specification Tests: If the empirical model provides a good approximation to the process that generates Y_{it}, it should fit the observed export market participation patterns up to a plant-specific time-invariant component plus serially-uncorrelated noise. Further, these disturbances should be orthogonal to each other and to the vector Z_{it}. As discussed by Andrews (1988) and Rust (1992), specification tests can be used to test these assumptions on the disturbance term jointly with our assumptions on the role Z_{it}, Y_{it-1} and $\tilde{Y}_{it-j}$.

Specifically, using the estimated parameters of equations (4) and (5), we will repeatedly combine actual series on the strictly exogenous variables (Z_{it}) with random draws on α_i, φ_{it} and ω_{it} to generate Y_{it} trajectories, plant by plant. If the model is valid, these simulated participation patterns for the years $J+1$ through T will differ from the actual ones only because of the random realizations on α_i, φ_{it} and ω_{it}. Thus each possible sequence of zeros and ones should occur with approximately the same frequency in both the simulated and actual data. Andrews' (1988) chi-square statistic provides a metric for comparing the two sets of frequencies.[24]

V. Econometric Results on Export Participation

In this section we report parameter estimates of equation (4). For all of the results reported here, we focus on four major exporting industries during the period 1981 through 1989. These industries are food products, textiles, paper products, and chemicals.[25] We limit our sample to the

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[24] Further details of the test are provided in the Appendix.

[25] We wish to limit the analysis to those industries in which Colombia appears able to compete in international markets. The industries were chosen because they account for a substantial percentage of total manufactured exports and have a relatively high proportion of plants participating in the export market. These four industries account for 58.6 percent of the value of manufactured exports in 1984 and 59.1 percent in 1987. The percentage of plants in our sample that export

650 plants in these industries that were in operation in each sample year.[26] This sample is not representative of the population of manufacturing plants, however, it is appropriate for examining the effects of sunk costs on established producers.[27] The final data set consists of 9 annual observations, covering the years 1981-1989, for each of 650 plants; a total of 5850 observations. The observations for 1981-1983 are treated as the three pre-sample years and are used to control for the initial-conditions problem using equation (5). The observations for 1984-1989 are used to estimate the role of sunk costs using equation (4).

The parameter estimates for equation (4) are reported in Table 3 for several model specifications. The most general model, reported in the first column, includes three lags of past participation (Y_{it-1}, $\tilde{Y}_{it-2}$, $\tilde{Y}_{it-3}$) as well as interaction terms involving Y_{it-1} and year dummy variables. These interaction terms allow the sunk costs of a new exporter to vary over time with market conditions.[28] The remaining three columns in Table 3 report results for models that restrict the

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averages 8.45 percent per year in the food industry, 19.7 in textiles, 15.4 in paper products, and 45.3 in chemicals.

[26] The main difference between the continuing group of plants we analyze and the plants that exit production over the period is that the latter group has a lower rate of entry into the export market, averaging 1.9 percent per year, and a lower degree of persistence once in, averaging .795. The differences, however, do not alter the general conclusion that transition rates, particularly for plants that do not export, are low and persistence is high. When examining the plants that entered production over the period, a pattern of export market transitions very similar to the plants we analyze is observed, particularly after the plants have been in operation for a few years. The main implication of these patterns for our analysis is that focusing solely on the group of continuing plants does not distort the patterns of export market transitions present in the manufacturing sector. It is this pattern of export market transitions that is important in estimating the econometric model.

[27] A more general framework would treat each plant as making simultaneous decisions to enter or exit production and enter or exit the export market. In this case, each plant could be viewed as choosing among four alternatives: do not produce, produce only for the domestic market, produce only for the export market, or produce for both. This approach is unnecessarily complicated for modeling the export decision in Colombia because there are no producers that sell only in the export market. In addition, very few plants enter production and the export market at the same time. As a result, focusing on the exporting behavior of plants that are already in operation, as we do, provides a reasonable starting point for analyzing the export determinants in Colombian manufacturing.

[28] We also estimated models that allowed sunk costs to vary with observable plant characteristics by including interactions between lagged participation and plant size, business type, and industry. None of these additional interactions were ever statistically significant and we do not report them here. We also generalized the pattern of time variation in sunk costs by interacting year dummies and the two period lag in participation. The additional coefficients from this specification were also insignificant.

coefficients on past participation. The second column restricts the interaction terms with Y_{it-1} to equal zero, implying sunk costs do not vary over time. The third column restricts the coefficents on $\tilde{Y}_{it-2}$ and $\tilde{Y}_{it-3}$ to equal zero, implying that the sunk costs of entry are the same for all non-exporting plants regardless of whether they had ever been exporters. The fourth column combines the previous two sets of restrictions. The final column restricts past participation to have no effect on current participation and is consistent with no sunk entry or exit costs.

Specification tests are reported at the bottom of each column in Table 3.[29] These chi-square statistics have 5 degrees of freedom, and provide an omnibus test of whether the model's parameterization and distributional assumptions are valid. The test does not reject the specification in either column 1 or 2 - both these models include three lagged values of past export participation. However, the models in columns 3, 4, and 5, which all restrict past participation to have at most one lag, are all rejected by the specification test. This implies that Y_{it} does not follow a first-order Markov process. Given that model 1 performs very well by this metric, we make it the focus of our discussion below.

Sunk Cost Parameters: Consider first the coefficients on Y_{it-1}, $\tilde{Y}_{it-2}$, $\tilde{Y}_{it-3}$, and Y_{it-1} interacted with time dummies. Together, these parameters isolate the importance of sunk costs. Using a likelihood ratio test to compare model 1 and model 5, we find that they are jointly significant with a $\chi^2(8)$ statistic of 206.06. This finding supports the basis premise of the hysteresis literature -- that there are substantial sunk costs involved in entering or exiting the export market.

[29] There are $2^6 = 64$ possible time paths for Y_{it} over the 1984-1989 period, some of which are quite rare. Accordingly, to improve the power of our test, we group these according to initial exporting status and the number of switches in export status over time, arriving at six types of trajectories. The categories of plant-specific Y_{it} trajectories are: all ones, all zeros, begin as a one and switch once, begin as a one and make multiple switches, begin as a zero and switch once, and begin as a zero and make multiple switches. The actual and predicted frequencies for these six trajectories are reported in the Appendix.

Looking at individual coefficients, we find that lagged export participation Y_{it-1} has a strong positive effect on the probability of exporting, as expected. Further, sunk entry costs vary significantly over time, as coefficients on the interaction terms between Y_{it-1} and the time dummies imply.[30] The sign pattern on the coefficients implies that the sunk costs fell from 1984 to 1985 for the typical plant, and rose steadily thereafter. One interpretation is that it is easier to break into an expanding world market than a shrinking one: 1985 was a year of relatively robust global expansion and overvaluation in the United States, while the late 1980s were a period of relatively slow growth in the United States.

As shown in equation (4), the coefficients on $\tilde{Y}_{it-2}$ and $\tilde{Y}_{it-3}$ measure the sunk costs of a new exporter minus the sunk costs of a plant that last exported two or three years earlier, respectively. They indicate that experience two years ago *does* significantly reduce sunk costs, relative to the costs of a plant that has never exported. That is, the benefits of past export market participation do not depreciate fully upon exit. On the other hand, the coefficent on $\tilde{Y}_{it-3}$ is not significantly different than zero, implying that plants that last exported three years earlier face re-entry costs roughly as large as those faced by a plant entering the market for the first time. Hence our choice of a three year lag structure appears to capture all of the relevant history.

Expected Profits from Exporting: The remaining coefficients in Table 3 summarize the influence of year effects and plant characteristics on the expected profitability of exporting, net of sunk entry costs ($\pi^*_{it} - F^0_{it}$). (Recall that this expression gives the net return from exporting for a plant with no prior foreign market experience.)

The time dummies indicate there is variation over time, but only the 1987 and 1989

[30] A likelihood ratio test for the joint significance of these interaction terms is rejected at the .05 significance level with a $\chi^2(5)$ statistic of 14.54.

coefficients are significantly different than zero.[31] Net of entry costs, the expected future profits from exporting were highest in 1986. This conforms to the relatively rapid entry rate between 1986 and 1987 (Table 2), and may have reflected producer anticipation of the coming years of currency depreciation. The incentives to begin exporting soon dissipate, however, and by 1989 expected net profits reach their in-sample low. Again, this is consistent with the falling entry rates in Table 2.

Interestingly, the decline in net expected profits appears to be due to rising sunk costs rather than falling expectations of gross operating profits. Relative to the base year, sunk entry costs were 1.08 higher in 1989, while expected profits net of sunk costs were only .847 lower. In fact, if exit costs (X_i) were zero, producers *already* exporting were doing better in 1989 than in any other year. One interpretation of this pattern is that the real exchange rate, which was favorable to exporters in 1989, determines expected operating profits for incumbents, while the strength of the world economy, which was ebbing in 1989, determines the ease with which new exporters can break in.

Net export profitability also varies systematically with observable plant characteristics. Notably we find that increases in plant size (measured by the plant's capital stock), increases in age, and corporate ownership all increase the probability of exporting. The plant size result may reflect scale economy-based exporting, as in Krugman (1984).[32] Alternatively, since efficient plants tend to grow relative to others, capital stock may simply be serving as a proxy for productivity. The age coefficient may also pick up cost differences among producers. If market forces select out inefficient producers then older plants will tend to be more competitive in world markets, either because of cost advantages that cannot be imitated by rivals or because they have had time to move down a learning

[31] The hypothesis that the time dummies are jointly equal to zero is rejected with a likelihood ratio test. The test statistic is 23.48 versus a $\chi^2(5)$ critical value of 15.1 at the .01 significance level.

[32] Combined with our finding that sunk costs do *not* rise with size, it is consistent with the well-known positive correlation between plant size and export participation. See Caves (1989) and Berry (1992) for a review of the evidence relating size and propensity to export.

curve.[33] Even if the annual pay-off from exporting were the same for young and old plants, the young ones would perceive less return to breaking into the market because they are less likely to survive.

Location matters as well, presumably because of transport costs. Bogota, the base category in the model, is land-locked in the Andes mountain range, and plants there are among the least likely to produce for foreign markets. Cali and Medellin are also inland, but less mountainous and closer to the coast. Nonetheless, we estimate that these cities are as unlikely to serve as a base for exporters as Bogota. Perhaps their locational advantage is offset by their lack of Bogota's agglomeration economies. Finally, the port cities of Cartagena and Baranquilla are most likely to host exporting plants.

Interestingly, neither wage rates nor export prices relative to domestic output prices are significant determinants of exporting behavior. This should *not* be interpreted to mean that prices don't matter; time dummies have already controlled for general movements in relative prices, and the plant-specific price variables therefore reflect across-plant deviations from average trends that can result from local market conditions, measurement error, and differences in input or output quality. Although we would expect increases in the export price to increase export market participation, we have no strong priors on whether changes in the cost of labor should make exporting more or less attractive than servicing the domestic market.

Unobserved Plant Heterogeneity and Noise: The final sources of variation in export status are unobserved error components: persistant plant heterogeneity, α_i, and transitory noise, ω_{it}. As shown in Table 3, .336 of the total unobserved variation is due to persistent heterogeneity, and this

[33] A decline in the probability of failure as a plant ages has been found by Roberts (1994) for Colombia and by Tybout (1994) for Chile. Liu and Tybout (1994) also find that failing plants in Colombia are systematically less productive than surviving plants. Both patterns have been found in data from the U.S.. (see Evans (1987), Dunne, Roberts, and Samuelson (1989), and Bailey, Hulten and Campbell (1992)).

fraction is significantly different than zero. Once this error component is controlled for, however, our specification tests indicate that remaining unobserved variation is serially uncorrelated and orthogonal to the set of explanatory variables.

In the pre-sample years, the fraction of variation due to unobserved plant effects is much higher (.928). This is simply because the lagged participation variables are not used to predict Y_{it} during 1981-1983, and their effect is shifted to the disturbance. An analogous effect is present in model 5, which leaves out lagged participation variables for all years. There, α_i accounts for .817 of total unexplained variation. Model 5 also demonstrates that lagged participation is strongly correlated with the vector Z_{it}. Note that without Y_{t-1}, $\tilde{Y}_{it-2}$, and $\tilde{Y}_{it-3}$, the remaining variables assume a larger role in predicting exporting status. All of these results confirm that export market participation equations without dynamics are seriously mis-specified.

Sunk Costs, Heterogeneity, and Export Probabilities: Table 4 quantifies the effects of observable plant characteristics, unobserved heterogeneity, and past participation on current export market participation. It is based on estimates of the unrestricted model reported in column 1 of Table 3. The three panels allow the observable determinants of export profitability (age, capital stock, industry etc.) to vary. Plants at the 25th, 50th, and 75th percentile of βZ_{it} are compared. Within each panel plants are distinguished by whether they never exported ($Y_{t-1} = \tilde{Y}_{it-2} = \tilde{Y}_{it-3} = 0$), last exported three years earlier ($Y_{t-1} = \tilde{Y}_{it-2} = 0$; $\tilde{Y}_{it-3} = 1$), last exported two years ago ($Y_{t-1} = \tilde{Y}_{it-3} = 0$; $\tilde{Y}_{it-2} = 1$), or exported last year ($Y_{t-1} = 1$; $\tilde{Y}_{it-2} = \tilde{Y}_{it-3} = 0$). The rows of the table summarize the effect of unobserved heterogeneity by allowing the normally-distributed permanent plant component α_i to vary from -2 to 2.

Export history matters for plants that have either above-average βZ_{it} or above-average α_i values. Thus, although only 25 percent of our sample ever exported, roughly 75 percent of the plants might well remain exporters if they were given some foreign market experience. (The probabilities that they would do so range from 14 percent to nearly 100 percent.) Expected profits from exporting for the remaining plants are so low that if they were somehow given export market experience it would not be sufficient to make them continue in the export market.

Table 4 also demonstrates how quickly experience depreciates. The difference in export probabilities between otherwise comparable plants that exported last year and those that last exported two years ago is substantial, ranging from .30 to .40. However, the difference between plants that last exported two years ago and plants that last exported three years ago is substantially smaller, ranging from .10 to .15, and plants that last exported three years ago are not much more likely to do so than those that never did.

Finally, inferences about the effects of changing macro conditions can also be drawn from Table 4.[34] For example, relative to 1984, the less favorable macro conditions that prevailed in 1989 shifted the probabilities that non-exporters would begin exporting approximately to those in the row directly above.

Overall, the results reported in Tables 3 and 4 reveal that the export participation decision is affected by time-period or macro conditions, observable plant cost or demand variables, unobserved time-invariant plant heterogeneity, and -- importantly for the sunk-cost hysteresis models -- prior export market experience. The latter has a particularly substantial effect on the probability a plant exports and this is consistent with the plant facing significant entry costs in the export market.

VI. Conclusions

In an attempt to explain the asymmetric patterns of export and import adjustment resulting from exchange rate movements, a recent group of papers develops theoretical models that rely on

[34] Since a one-unit change in α corresponds to a .71 shift in b_k (by equation 6 and the definition of ξ_{it}, anything that shifts b_k by .71 corresponds to a one-row upward movement in Table 4.

sunk costs of entry to produce hysteresis in trade flows. Beginning with the assumption that entry into foreign markets requires exporters to incur some sunk costs, these models demonstrate that temporary changes in market variables, such as exchange rates, can result in permanent changes in market structure and exports because of the entry and exit of exporters. While the hypothesized response in these models is clearly a micro process of entry and exit, empirical tests to date have relied on aggregate or sectoral data on trade flows and prices.

In this paper we develop an econometric model of a plant's decision to export and use it to test one of the key assumptions underlying the theoretical models of sunk cost hysteresis. We utilize micro data for a large group of manufacturing plants in Colombia from 1981-1989 and directly examine the determinants of a plant's export decision for consistency with the theory. The implication of the sunk cost models that we test is that past participation in the export market will have a significant effect on the probability of exporting in the current period. Equivalently, the presence of sunk entry or exit costs will lead to true state dependence in the export decision.

The econometric results, which control for both observed and unobserved sources of plant heterogeneity, indicate that prior export participation has a significant effect on the probability a plant exports. Equivalently, we reject the hypothesis that sunk costs are zero. The empirical results also reveal that the re-entry costs of plants that have been out of the export market for a year are substantially less than the costs of a first-time exporter. Beyond a one year absence, however, the re-entry costs are not significantly different than those faced by a new exporter. This is consistent with the view that an important source of sunk entry costs for Colombian exporters is the need to accumulate information on demand sources, information that is likely to depreciate upon exit from the market.

While the results indicate that sunk costs are a significant source of export market persistence, both observed and unobserved plant characteristics also contribute to an individual plant's export behavior. Plants that are large, old, and owned by corporations are all more likely to export. Variation in unobserved sources of difference in profitability can lead to as much as a 30 percentage point difference in the probability of exporting for a plant with no prior experience.

This combination of plant heterogeneity and sunk costs implies that the response of aggregate or sectoral exports to changes in policy or the macro environment will likely be idiosyncratic with respect to country and time period. The magnitude of the supply response will depend upon the number and type of plants already participating in the export market, the stability or permanence of the policy regime, the magnitude of temporary shocks, and the sunk costs of entering a new market. The latter, in turn, is likely to vary with the degree of information producers have about foreign markets, the type of market they are likely to enter, the type of product being exported, and the policy regime. Given the number of idiosyncratic forces at work, it is not surprising that standard empirical export supply functions have exhibited marked instability across countries and time.

Finally, our findings suggest that countries undertaking export promotion policies should distinguish measures aimed at expanding the export volume of exisiting exporters from policies aimed at promoting the entry of new exporters. The latter include actions directed at reducing entry costs and uncertainty, such as providing information about potential markets, developing exporting infrastructure, or providing a stable macro and policy environment. If entering the export market is a more significant hurdle for firms than expanding their output once in the market, these entry promotion policies may be more effective at expanding exports than direct subsidies based on the value of exports.

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Table 1

Colombian Manufactured Exports 1981-1989
(nineteen three-digit SITC industries)

	1981	1982	1983	1984	1985	1986	1987	1988	1989
Real Effective Exchange Rate Index	117.8	125.7	125.1	114.6	100	74.4	66.3	64.4	62.7
Real Value of Exports	109.6	110.8	113.7	95.9	110.8	151.2	169.4	189.1	276.2
Export Subsidy Rate	.055	.055	.066	.099	.092	.047	.047	.042	.044
Number of Exporting Plants	667	676	615	585	653	705	707	735	816
Proportion of Plants that Export	.129	.128	.113	.107	.117	.122	.119	.124	.135

Table 2

Plant Transition Rates in the Export Market 1982-1989

Year t Status	Year t+1 Status	1982-1983	1983-1984	1984-1985	1985-1986	1986-1987	1987-1988	1988-1989	Average 1982-1989
		(Nineteen three-digit manufacturing industries)							
No Exports	No Exports	.974	.971	.957	.963	.973	.972	.958	.967
	Exports	.026	.029	.043	.037	.026	.028	.042	.033
Exports	No Exports	.168	.135	.131	.108	.158	.086	.107	.128
	Exports	.832	.865	.869	.892	.842	.914	.893	.872
		(Four major exporting industries)							
No Exports	No Exports	.971	.969	.972	.960	.983	.972	.985	.973
	Exports	.029	.031	.028	.040	.017	.028	.015	.027
Exports	No Exports	.108	.101	.152	.124	.149	.085	.054	.110
	Exports	.892	.899	.848	.876	.851	.915	.946	.890

Table 3
Dynamic probit model of export participation
(standard errors in parenthesis)

Explanatory variable	*Model 1*	*Model 2*	*Model 3*	*Model 4*	*Model 5*
Intercept	-8.312*(1.456)	-8.393*(1.426)	-9.285*(1.433)	-9.381*(1.458)	-16.954*(1.710)
Y_{t-1}	1.933*(.261)	2.170*(.152)	1.707*(.245)	1.865*(.124)	
Y_{t-1}*(1985 Dummy)	-.106 (.313)		-.170 (.315)		
Y_{t-1}*(1986 Dummy)	-.009 (.318)		-.051 (.321)		
Y_{t-1}*(1987 Dummy)	.185 (.324)		.202 (.327)		
Y_{t-1}*(1988 Dummy)	.261 (.328)		.296 (.334)		
Y_{t-1}*(1989 Dummy)	1.082*(.368)		1.112*(.371)		
$\tilde{Y}_{t-2}$	.732*(.195)	.777*(.193)			
$\tilde{Y}_{t-3}$	.347 (.248)	.340 (.244)			
1985 Dummy	-.155 (.201)	-.215 (.161)	-.130 (.197)	-.221 (.163)	-.276 (.178)
1986 Dummy	.004 (.189)	.018 (.159)	.026 (.187)	.019 (.161)	-.068 (.179)
1987 Dummy	-.469*(.220)	-.398*(.169)	-.498*(.218)	-.436*(.171)	-.476*(.189)
1988 Dummy	-.248 (.202)	-.149 (.170)	-.286 (.201)	-.205 (.174)	-.330*(.193)
1989 Dummy	-.847*(.253)	-.378*(.178)	-.902*(.249)	-.458*(.182)	-.520*(.198)
ln($Wage_{t-1}$)	.216 (.147)	.231 (.145)	.256 (.149)	.236 (.150)	.470*(.173)
ln(Export $price_{t-1}$)	-.062 (.062)	-.059 (.061)	-.064 (.062)	-.084 (.067)	.014*(.083)
ln(K_{t-1})	.247*(.045)	.239*(.043)	.291*(.041)	.288*(.043)	.537*(.052)
Age_{t-1}	.337*(.115)	.330*(.114)	.379*(.115)	.427*(.116)	.701*(.169)
Corporation	.368*(.158)	.333*(.155)	.365*(.154)	.450*(.161)	.396 (.229)
Textiles Ind. Dummy	1.018*(.198)	.999*(.194)	1.124*(.179)	1.221*(.196)	2.392*(.259)
Paper Ind. Dummy	.251 (.183)	.254 (.180)	.250 (.189)	.237 (.190)	.940*(.273)
Chemicals Ind. Dummy	.532*(.179)	.503*(.176)	.875*(.187)	.628*(.181)	2.996*(.285)
Cali/Medellin	-.099 (.143)	-.099 (.139)	-.060 (.140)	-.123 (.143)	.547*(.190)
Other region	.378*(.139)	.371*(.136)	.409*(.136)	.456*(.141)	1.430*(.204)
Var(α)	.336*(.076)	.314*(.076)	.416*(.053)	.446*(.059)	.817*(.016)
Cov(φ,α)	.928*(.013)	.928*(.013)	.917*(.013)	.926*(.012)	.779*(.025)
ln(L)	-845.80	-853.07	-854.09	-860.86	-948.83
Specification test	5.121	5.519	12.065*	9.533**	55.112*

* Reject null hypothesis at the .05 significance level.

** Reject null hypothesis at the .10 significance level.

Table 4
Predicted probability of exporting

Plant effect (α)		*25th percentile of βZ_{it}*				*50th percentile of βZ_{it}*				*75th percentile of βZ_{it}*			
	Y_{t-1} =	0	0	0	1	0	0	0	1	0	0	0	1
	$\tilde{Y}_{t-2}$ =	0	0	1	0	0	0	1	0	0	0	1	0
	$\tilde{Y}_{t-3}$ =	0	1	0	0	0	1	0	0	0	1	0	0
-2		.000	.000	.000	.006	.000	.000	.001	.024	.001	.002	.006	.096
-1		.000	.000	.001	.037	.001	.002	.007	.104	.006	.015	.037	.277
0		.001	.004	.011	.141	.007	.016	.040	.291	.035	.071	.140	.547
1		.011	.026	.059	.358	.038	.077	.149	.564	.135	.225	.356	.797
2		.056	.108	.197	.636	.144	.238	.371	.808	.348	.482	.633	.938

Appendix: Specification Tests

Andrews' (1988) specification test compares realized (Y_{it}, Z_{it}) trajectories in our sample to expected trajectories based on the estimated model. To construct his χ^2 statistic, we first partition the possible (Y_{it}, Z_{it}) trajectories into a limited number of cells. In our case we have $2^6 = 64$ possible trajectories for Y_{it}, some of which are very unusual. To avoid cells that are nearly empty we distinguish six types of trajectories: $Y_{it} = 0 \ \forall \ t$; $Y_{it} = 0$ initially, but switches *once* during the sample period; $Y_{it} = 0$ intially and switches *at least twice* during the sample period; $Y_{it} = 1 \ \forall \ t$; $Y_{it} = 1$ initially, but switches *once* during the sample period; and $Y_{it} = 0$ initially and switches *at least twice* during the sample period. Letting the vector indicator function $\Psi(Y_{i,t+1}, Y_{i,t+2}, \ldots, Y_{i,t+T})$ map Y_{it} sequences into these 6 cells, the observed frequencies of the different trajectories in our sample is the vector $P_n(\Psi) = \frac{1}{n}\sum_{i=1}^{n} \Psi(Y_{i,t+1}, Y_{i,t+2}, \ldots Y_{iT})$. Elements of this vector are reported in Table A1.1 below.

Table A1.1: Observed versus Predicted Frequencies of Y_{it} Trajectories (based on Column 1, Table 3)

Trajectory type	Observed Frequencies	Expected Frequencies
always a non-exporter	.761	.743
begin as a non-exporter, switch once	.046	.051
begin as a non-exporter, switch at least twice	.045	.052
always an exporter	.098	.089
begin as an exporter, switch once	.017	.028
begin as an exporter, switch at least twice	.032	.037

Next, to generate model-based expected values for each of these cells, we use estimated parameter values from Table 3 in conjunction with the observed Z_{it} trajectories and random draws on α_i, φ_{it} and ω_{it} to repeatedly simulate Y_{it} sequences, plant by plant. (The reported tests are based on 200 simulations per plant.) Distributions for each of these random variables are based on the assumptions described in section III. Averaging over all of the outcomes for the i^{th} plant, we get the probabilities it will fall in each cell under the null hypothesis that our specification is correct: $Q(\Psi, Z_{iJ}, Z_{iJ+1}, \ldots, Z_{iT} \mid \mu, \gamma, \sigma_\alpha, \rho)$. Finally, averaging these probabilities over all plants, we obtain the expected sample-wide frequencies of each cell in the partition:

$$Q_n(\Psi) = \frac{1}{n} \sum_{i=1}^{n} Q(\Psi, Z_{iJ}, Z_{iJ+1}, \ldots, Z_{iT} \mid \mu, \gamma, \sigma_\alpha, \rho)$$

The expected frequencies generated by the model in column 1 of Table 3 are reported in the second column of Table A1.1. The test statistic is calculated as a quadratic form in the difference between the two columns, $X^2_{(K-1)} = (P_n - Q_n)'\hat{W}(P_n - Q_n)$; where the weighting matrix $\hat{W}$ is given by equation (15) in Andrews' appendix.

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