The Decision to Export in Colombia: An Empirical Model of Entry with Sunk Costs

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Recent theoretical models of entry predict that, in the presence of sunk costs, current market participation is affected by prior experience. This paper quantifies the effect of prior exporting experience on the decisions of Colombian manufacturing plants to participate in foreign markets. It develops a dynamic discrete-choice model of exporting behavior that separates the roles of profit heterogeneity and sunk entry costs in explaining plants' exporting status. Sunk costs are found to be significant, and prior export experience is shown to increase the probability of exporting by as much as 60 percentage points. (*JEL* F10, L10, C25)

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 series of papers Richard Baldwin, Avinash Dixit, and Paul Krugman have proposed an answer.[1] They begin from the assumption that nonexporters 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 conditions, 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 focused on asymmetries in the response of trade flows to exchange rate appreciation versus

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

[2] In their review of empirical studies of price and income elasticities for traded goods, Morris Goldstein and Mohsin S. Khan (1985 pp. 1087–92) 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.

depreciation.[3] One problem with empirical tests at this level is that a number of forces including changes in expectations, adjustment costs, and pricing-to-market behavior that have nothing to do with sunk costs can generate apparent asymmetries. Another limitation is that aggregate trade flow data cannot distinguish changes in the number of exporters from changes in the volume of shipments of continuing exporters and thus cannot be used to predict entry and exit patterns under different market conditions.[4] For these reasons, the central tenet of the hysteresis literature that sunk costs affect foreign-market participation patterns has not been directly addressed empirically.

In this paper we test for sunk-cost hysteresis by directly analyzing entry and exit patterns in plant-level panel data. To do so, we develop and estimate a dynamic discrete-choice model that expresses each plant's current exporting status as a function of its previous exporting experience, observable characteristics that affect its future profits from exporting, and unobserved serially correlated shocks. The conditional effect of a plant's exporting history on its current exporting status allows us to infer the importance of sunk costs. The model also allows us to quantify the effects of producer characteristics, macro shocks, and prior experience on the probability of participating in the market. With minor modifications our framework can be used to analyze a wide range of firm diversification or investment activities including the decision to enter a new geographic market, introduce a new product, initiate an advertising campaign, or undertake capital or R&D investments.

The data we use describe the export patterns of 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 reject the hypothesis that sunk costs are zero. This implies that prior export-market experience significantly affects the current decision to export, and the policy implications stressed in the hysteresis literature are empirically relevant. Further, although experience in foreign markets is important, its effect depreciates fairly quickly over time. A plant that exported in the prior year is up to 60 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.

In the next section of the paper we provide an overview of the theoretical sunk-cost model. Section II summarizes the patterns of export participation among Colombian manufacturing plants between 1981 and 1989. Sec-

[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, Joseph 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, Robert M. 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. David C. Parsley and Shang-Jin 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] José Manuel Campa (1993) does look at the correlates of foreign-firm investments in the 61 U.S. wholesale trade industries over the 1981–1987 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 United States. He also reports that an industry's sunk costs, which are proxied by the advertising–sales ratio and the ratio of fixed assets to net worth of firms in the industry, are negatively correlated with foreign-firm entry. Both findings are consistent with the hysteresis model. Although not in a trade context, related work by Timothy F. Bresnahan and Peter C. 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, persistence in behavior due to permanent differences in profitability across producers can create the appearance of sunk costs in their model.

tion III develops our econometric model of the export decision, and Section IV presents our results. We summarize and draw conclusions in Section V.

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

To motivate our empirical work, we begin by reviewing the basic theory (see footnote 1 for references). For each period t, let the ith plant's expected gross profits when exporting differ from its expected gross profits when not exporting by the amount $\pi_i(\mathbf{p}_t, \mathbf{s}_{it})$. Here $\mathbf{p}_t$ is a vector of market-level forcing variables that the plant takes as exogenous (e.g., the exchange rate and foreign demand conditions), and $\mathbf{s}_{it}$ 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 $\pi_i(\mathbf{p}_t, \mathbf{s}_{it})$ represents the increment to expected profits associated with exporting in year t, assuming that the profit-maximizing level of exports is always chosen.

These gross profits have not been adjusted for the sunk costs of foreign-market entry or exit. Assume that if the ith plant last exported in year $t - j$ ($j \geq 2$) it faces a reentry cost of F_i^j, so upon resuming exports in year t it earns $\pi_i(\mathbf{p}_t, \mathbf{s}_{it}) - F_i^j$. Similarly, if the plant had never exported previously, it faces an entry cost of F_i^0 and earns $\pi_i(\mathbf{p}_t, \mathbf{s}_{it}) - F_i^0$ in its first year exporting. Finally, a plant that exported in period $t - 1$ earns $\pi_i(\mathbf{p}_t, \mathbf{s}_{it})$ 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 reentry costs to depend on the length of absence from the market. This could reflect the decreasing relevance 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 $\mathbf{Y}_{it}^{(-)} = \{Y_{i,t-j} \mid j = 0, \dots, J_i\}$, where J_i is the age of the plant. Then period-t exporting profits are

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

where $\tilde{Y}_{i,t-j} = (Y_{i,t-j}\prod_{k=1}^{j-1}(1 - Y_{i,t-k}))$. This last expression summarizes the plant's most recent exporting experience: $\tilde{Y}_{i,t-j} = 1$ if the plant was last in the export market j years earlier and 0 otherwise.

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

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

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

$$V_{it}(\Omega_{it}) = \max_{\mathbf{Y}_{it}}(R_{it}(\mathbf{Y}_{it}^{(-)}) + \delta E_t\{V_{i,t+1}(\Omega_{i,t+1}) \mid \mathbf{Y}_{it}^{(-)}\})$$

where E_t denotes expected values conditioned on the information set Ω_{it}. From the right-hand

[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.

side of this expression, it follows that the ith plant will be in the export market during period t if

$$(1)\quad \pi_i(\mathbf{p}_t, \mathbf{s}_{it}) + \delta[E_t(V_{i,t+1}(\Omega_{i,t+1}) \mid Y_{it}=1) - E_t(V_{i,t+1}(\Omega_{i,t+1}) \mid Y_{it}=0)] \geq F_i^0 - (F_i^0 + X_i)Y_{i,t-1} + \sum_{j=2}^{J_i} (F_i^0 - F_i^j)\tilde{Y}_{i,t-j}$$

where $-(F_i^0 + X_i)$ 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 III. It has several empirical implications which we will pursue. First, if there are no sunk costs, the participation condition collapses to $\pi_i(\mathbf{p}_t, \mathbf{s}_{it}) \geq 0$. 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. Hence the magnitude of sunk costs and the rate at which past experience depreciates can be identified.[6] Finally, this equation indicates that realizations on the variables $\mathbf{p}_t$ and $\mathbf{s}_{it}$ influence export decisions through their effect on $\pi_i(\mathbf{p}_t, \mathbf{s}_{it})$ *and* their effect on the expected future value from 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.

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

II. The Pattern of Export Participation in Colombia

Before discussing estimation issues and results, we discuss our data base, review the export environment in Colombia, and provide some aggregate evidence on the pattern of export-market participation during the 1980's.

A. *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.

B. *The Policy Regime and Export Participation Rates*

Table 1 illustrates the basic patterns over the sample period for the 19 major exporting industries in the Colombian manufacturing sector.[8] In general terms, the macroeconomic environment in Colombia was not conducive to

[7] The census data and matching process are discussed in greater detail in Roberts (1996).

[8] The 19 industries and their ISIC 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), nonmetal 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.

TABLE 1—COLOMBIAN MANUFACTURED EXPORTS 1981–1989 (19 THREE-DIGIT ISIC INDUSTRIES)

Variable	1981	1982	1983	1984	1985	1986	1987	1988	1989
Real effective exchange rate index (1975 = 100)[a]	84.0	79.5	80.5	89.8	102.2	113.6	113.7	112.3	115.3
Quantity of exports (1986 = 100)[b]	58.5	64.1	63.8	59.1	66.4	100.0	88.6	103.1	127.2
Export subsidy rate	0.055	0.055	0.066	0.099	0.092	0.047	0.047	0.042	0.044
Number of exporting plants	667	676	615	585	653	705	707	735	816
Proportion of plants that export	0.129	0.128	0.113	0.107	0.117	0.112	0.119	0.124	0.135

[a]Source: José Ocampo and Leonardo Villar (1995). An increase in this variable corresponds to a devaluation of the Colombian peso.

[b]Source: Roberts et al. (1995). Figures describe export-oriented industries only.

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 1982. As shown in Table 1, this pattern was reversed after 1982, with the currency losing 43 percent of its value by 1986 before stabilizing. This partly reflected central-bank currency-market interventions to ease competitive pressures on tradable-goods producers.

The time-series pattern of manufactured exports roughly mirrors this movement in the exchange rate. The index of real manufactured exports from the major exporting industries was stable through 1984 and then increased rapidly through 1986, growing at an average annual rate of 35 percent over that period. Exports again grew rapidly between 1988 and 1989.

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 policies made it more difficult to import raw materials or capital goods which may have been necessary to increase the quality of manufactured products.[9] 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 declined thereafter.[10] 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 4-percent real appreciation between 1981 and 1983 was accompanied by a decline in the export participation rate from 0.129 to 0.113, but a much larger (28 percent) depreciation between 1984 and 1989 served only to increase the participation rate to 0.135. From this short time series it

[9] 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).

[10] 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.

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.[11] First, to sell in developed-country markets, Colombian producers were often required to invest in product-quality upgrading. Second, there was little exporting infrastructure in the form of trading companies or distribution agents. These companies typically provide transportation, customs, and shipping services, as well as information on 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 was 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 forgo 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, lack of 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.

C. *Entry and Exit in the Export Market*

The analytical model reviewed in Section I 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 the first column to the status in the second column. 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 $t+1$.[12] 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 mar-

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

[12] The data correspond to the group of plants that were in operation in each year during 1981–1989. There are 2,369 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.

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
A. *Nineteen three-digit manufacturing industries:*									
No exports	No exports	0.974	0.971	0.957	0.963	0.973	0.972	0.958	0.967
	Exports	0.026	0.029	0.043	0.037	0.026	0.028	0.042	0.033
Exports	No exports	0.168	0.135	0.131	0.108	0.158	0.086	0.107	0.128
	Exports	0.832	0.865	0.869	0.892	0.842	0.914	0.893	0.872
B. *Four major exporting industries:*									
No exports	No exports	0.971	0.969	0.972	0.960	0.983	0.972	0.985	0.973
	Exports	0.029	0.031	0.028	0.040	0.017	0.028	0.015	0.027
Exports	No exports	0.108	0.101	0.152	0.124	0.149	0.085	0.054	0.110
	Exports	0.892	0.899	0.848	0.876	0.851	0.915	0.946	0.890

ket, 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, of the plants that exported at some time during the 1981–1989 period, only 36 percent remained exporters for the whole sample period, and among plants that *did* change exporting status, 60 percent did so more than once.

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 differences across plants in the gross profit from exporting, $\pi_i(\mathbf{p}_t, \mathbf{s}_{it})$, 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.

III. An Empirical Model of Export-Market Participation

A. *The Estimating Equation*

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

$$\pi_{it}^* = \pi_i(\mathbf{p}_t, \mathbf{s}_{it}) + \delta[E_t(V_{it}(\Omega_{i,t+1}) \mid Y_{it} = 1) - E_t(V_{it}(\Omega_{i,t+1}) \mid Y_{it} = 0)]$$

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:

$$(2) \quad Y_{it} = \begin{cases} 1 & \text{if } \pi_{it}^* - F_i^0 + (F_i^0 + X_i)Y_{i,t-1} \\ & \quad + \sum_{j=2}^{J_t} (F_i^0 - F_i^j)\tilde{Y}_{i,t-j} \geq 0 \\ 0 & \text{otherwise.} \end{cases}$$

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 $\mathbf{s}_{it}$ and $\mathbf{p}_t$.[13] Alternatively, we could forgo identification of structural parameters, and approximate $\pi_{it}^* - F_i^0$ as a reduced-form expression in exogenous plant and market characteristics that are observable to producers in period t. The advantage of the

[13] James J. Heckman (1981b) provides detailed discussion of estimation of dynamic discrete-choice models. Zvi Eckstein and Kenneth I. Wolpin (1989) and John Rust (1997) summarize the literature on estimating structural dynamic models of discrete choice. Yacine Ait-Sahalia (1994) uses a structural-hysteresis model to impute sunk costs and fixed costs under the assumption of one domestic and one foreign producer per industry.

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 Jth-order Markov process, conditioned on exogenous variables. 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 exporting decisions to $\mathbf{s}_{it}$ and $\mathbf{p}_t$, we pursue the reduced-form approach.

To parameterize the reduced-form model, we assume that variation in $\pi_{it}^* - F_i^0$ 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}:

$$(3) \qquad \pi_{it}^* - F_i^0 = \mu_t + \boldsymbol{\beta}\mathbf{Z}_{it} + \varepsilon_{it}.$$

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 $\mathbf{p}_t$ in the analytical model. The vector $\mathbf{Z}_{it}$ controls for factors represented by $\mathbf{s}_{it}$ and F_i^0 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 ISIC level, a dummy variable to control for the ownership structure of the plant (proprietorship and partnership vs. corporation), and two locational dummies to distinguish the inland Bogotá and Medellín/Cali regions from the coastal cities.[14] The vector $\mathbf{Z}_{it}$ 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.[15] Relative prices and wages affect the attractiveness of domestic versus foreign markets. Capital stock and age proxy for efficiency differences: in addition to scale effects, studies of industrial evolution 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. For our basic specification we assume that all plants that have not exported for at least J years face the same entry costs, $F_i^0 = F^0$, and all plants that have not exported for $j < J$ years face the same sunk entry costs, $F_i^j = F^j$. (The assumption that entry costs are the same for all $j > J$ is not restrictive if a generous value for J is chosen.) Further, all plants currently exporting face the same exit cost, $X_i = X$. Then defining the parameters $\gamma^j = F^0 - F^j$ ($j = 2, \dots, J$) and $\gamma^0 = F^0 + X$ and substituting (3) into (2), we obtain our basic estimating equation:

$$(4) \quad Y_{it} = \begin{cases} 1 & \text{if } 0 \le \mu_t + \boldsymbol{\beta}\mathbf{Z}_{it} + \gamma^0 Y_{i,t-1} \\ & \qquad + \sum_{j=2}^{J} \gamma^j \tilde{Y}_{i,t-j} + \varepsilon_{it} \\ 0 & \text{otherwise.} \end{cases}$$

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^0 and the 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 by comparing the magnitude of the γ coefficients. By also including interaction terms between the lagged participation variables and plant characteristics or macro variables, we can test our assumption that sunk costs do not vary across plants or time. Finally, we can use equation

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

[15] 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.

(4) to study the importance of temporal (μ_t) and cross-plant ($\boldsymbol{\beta}\mathbf{Z}_{it}$) variation in net expected profits from exporting ($\pi^*_{it} - F^0_i$). In particular, we can impute probabilities of entry or exit in response to a given shift in exogenous variables for plants with different characteristics.

B. *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 $\mathbf{Z}_{it}$ 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 the persistence it induces in exporting status to sunk costs.[16]

We allow for two sources of serial correlation in ε_{it} by assuming it is the sum of a permanent, plant-specific component and a transitory, first-order autoregressive component: $\varepsilon_{it} = \alpha_i + \omega_{it}$, where $\omega_{it} = \rho\omega_{i,t-1} + \eta_{it}$. Here α_i is independently and identically distributed normal across plants, η_{it} is independently and identically distributed normal across plants and time, $\mathrm{Cov}(\mathbf{Z}_{it}, \varepsilon_{it}) = \mathrm{Cov}(\alpha_i, \omega_{it}) = 0 \quad \forall\, i, t$, and we normalize $\mathrm{Var}(\varepsilon_{it}) = 1$. The plant effect α_i represents unobservable differences in managerial efficiency, foreign contacts, and other factors that induce persistent plant-specific differences in the returns from exporting.[17] Transitory unobserved shocks to $\pi^*_{it} - F^0_i$ are represented by ω_{it}. This specification implies that $\mathrm{Cov}(\varepsilon_{it}, \varepsilon_{i,t-k})$ depends on two parameters: the fraction of the variance of ε_{it} that arises from the permanent component in the error, $\mathrm{Var}(\alpha)$, and the serial correlation in transitory shocks to exporting profits, ρ.

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 decision in years $J + 1$ through T. But $Y_{i,t-1}$ and $\tilde{Y}_{i,t-j}$ values corresponding to the first J years cannot be treated as exogenous determinants of Y_{it} ($J + 1 \le t \le 2J$) because each depends on α_i and previous realizations of ω, both of which are correlated with ε_{it}. Heckman (1981c) suggests dealing with this "initial-conditions" problem by using an approximate representation for Y_{it} when $t \le J$ 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 the J presample years can be represented with the equation

$$\pi^*_{it} - F^0_i = \boldsymbol{\lambda}\mathbf{Z}^{\mathrm{p}}_{it} + \varepsilon^{\mathrm{p}}_{it}$$

where $\mathbf{Z}^{\mathrm{p}}_{it}$ is a distributed lag in presample realizations on exogenous variables.[18] Then presample export-market participation is described by

$$(5) \qquad Y_{it} = \begin{cases} 1 & \text{if } 0 \le \boldsymbol{\lambda}\mathbf{Z}^{\mathrm{p}}_{it} + \varepsilon^{\mathrm{p}}_{it} \\ 0 & \text{otherwise} \end{cases}$$

instead of equation (4).

[16] This is the problem of "spurious state-dependence" discussed in the empirical literature on labor-market participation (see e.g., Heckman, 1981a, b).

[17] We do not control for the α_i by using plant-specific dummy variables because of the "incidental-parameters problem" discussed in J. Neyman and E. Scott (1948), Gary Chamberlain (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 $n \to \infty$ 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 or the model is dynamic, the bias can be substantial. See B. D. Wright and G. Douglas (1975), Heckman (1981c), and Cheng Hsiao (1986 pp. 159–61) 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 $T = 8$.

[18] In the empirical work we include all of the plant characteristics in $\mathbf{Z}_{it}$ described above as explanatory variables in $\mathbf{Z}^{\mathrm{p}}_{it}$. Also included are two-year lagged values of the plant's wages, capital stock, and export price.

As with the error ε_{it} in equation (4), we assume that the presample disturbance term has the form $\varepsilon_{it}^{\mathrm{p}} = \alpha_i^{\mathrm{p}} + \omega_{it}^{\mathrm{p}}$, where $\omega_{it}^{\mathrm{p}} = \rho^{\mathrm{p}}\omega_{i,t-1}^{\mathrm{p}} + \eta_{it}^{\mathrm{p}}$. Here α_i^{p} is independently and identically distributed normal across plants, η_{it}^{p} is independently and identically distributed normal across plants and time, $\mathrm{Cov}(\mathbf{Z}_{it}^{\mathrm{p}}, \varepsilon_{it}^{\mathrm{p}}) = \mathrm{Cov}(\alpha_i^{\mathrm{p}}, \omega_{it}^{\mathrm{p}}) = 0 \quad \forall\, i, t$, and we normalize $\mathrm{Var}(\varepsilon_{it}^{\mathrm{p}}) = 1$. Critically, we allow the plant effects α_i and α_i^{p} to be correlated, and we allow serial correlation in transitory noise between periods J and $J + 1$. Thus we allow our model to attribute correlation between Y_{it} and the lagged participation variables, $Y_{i,t-1}$ and $\tilde{Y}_{i,t-j}$ $(j \leq J)$ to serial correlation in disturbances, even in the early sample years when $J + 1 < t < 2J$. This initial-conditions correction adds the coefficient vector $\boldsymbol{\lambda}$, the variance component $\mathrm{Var}(\alpha_i^{\mathrm{p}})$, the serial correlation parameter ρ^{p}, and $\mathrm{Corr}(\alpha_i, \alpha_i^{\mathrm{p}})$ to the model's parameters.[19]

Given that we model the disturbances ε_{it} and $\varepsilon_{it}^{\mathrm{p}}$ as random effects plus AR(1) processes, the system of T participation equations (4) and (5) identifies parameters using both cross-sectional and temporal variation in the data. 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 economywide fluctuations in macroeconomic conditions, and to unobserved plant-specific shocks, which combine with changes in $\mathbf{Z}_{it}$ to induce temporal variation in Y_{it}. Note that, even if there were no variation in $\mathbf{Z}_{it}$, the transitory shocks (ω_{it}^{p} and ω_{it}) would suffice to identify the coefficients on $Y_{i,t-1}$ and $\tilde{Y}_{i,t-j}$.

In principle, our system could be estimated using maximum-likelihood (ML) techniques, but that would involve T-dimensional integrals. Accordingly, we use Michael P. Keane's (1994) estimator which adapts the method of simulated moments (MSM) developed by Daniel McFadden (1989) and Ariel Pakes and David Pollard (1989) to panel data. In essence, Keane's technique involves choosing a trial set of parameter values, taking random draws from the implied multivariate distributions that characterize the error terms, and combining them with actual trajectories of the exogenous variables ($\mathbf{Z}_{it}$) to simulate probabilities of Y_{it} trajectories for each plant. Differences between observed and expected Y_{it} trajectories for each plant are aggregated using method-of-moments weights to obtain a metric for goodness of fit, and the parameter vector is varied until this metric is minimized. Keane's estimator is consistent and asymptotically efficient in the number of simulations.[20]

If we impose the constraint $\rho = \rho^{\mathrm{p}} = 0$, the transitory errors ε_{it} and $\varepsilon_{it}^{\mathrm{p}}$ are serially uncorrelated, and the problem of T-dimensional integration does not arise. Thus, under this maintained hypothesis, ML estimation using bivariate Gaussian quadrature to integrate out α_i and α_i^{p} becomes feasible (J. S. Butler and Robert Moffitt, 1982), and we can avoid the efficiency loss associated with simulation estimators. In addition, as outlined in the Appendix, Donald W. K. Andrews's (1988) omnibus specification test is feasible when we use ML estimation. For these reasons we report both MSM and ML estimates for versions of the model that restrict $\rho = \rho^{\mathrm{p}} = 0$.

IV. Econometric Results on Export Participation

To estimate the econometric model in equations (4) and (5) we focus on four major exporting industries (food products, textiles, paper products, and chemicals) during the period 1981–1989.[21] We limit our sample to the

[19] Although equation (5) is an imperfect representation of the process generating the data, simulation evidence suggests that Heckman's procedure handles the initial-conditions problem reasonably well. 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 presample Y_{it} realizations as functions of data and estimable parameters.

[20] The technique reduces dimensionality problems by simulating transition probabilities rather than probabilities for entire sequences of Y_{it} realizations. These simulations are done using the highly accurate Geweke-Hajivassiliou-Keane algorithm. Details are found in Keane (1994). A program that implements Keane's estimator was kindly provided to us by David Ribar.

[21] 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

650 plants in these industries that were in operation in each sample year.[22] 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.[23] The final data set consists of nine annual observations, covering the years 1981–1989, for each of 650 plants, a total of 5,850 observations. The observations for 1981–1983 are treated as the $J = 3$ presample years and are used to control for the initial-conditions problem using equation (5). The observations for 1984–1989 are used to estimate the parameters of interest in equation (4).

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 ($Y_{i,t-1}$, $\tilde{Y}_{i,t-2}$, $\tilde{Y}_{i,t-3}$) and allows for serially correlated transitory errors.[24] The second and third columns report MSM and ML estimates of the model that restricts $\rho = 0$ so that there is no serial correlation in the transitory errors. The fourth and fifth columns report MSM and ML estimates of a model that also restricts the relevant exporting history to just a single lagged value of the participation variable.

A. *Sunk-Cost Parameters*

Consider first the coefficients on $Y_{i,t-1}$, $\tilde{Y}_{i,t-2}$, and $\tilde{Y}_{i,t-3}$ in the most general specification, reported in column (i). Together, these parameters isolate the importance of sunk costs. Using a Wald test, we reject the hypothesis that the three coefficients are jointly equal to zero, with a $\chi^2_{[3]}$ statistic of 11.53. Thus, even after controlling for a general form of serial correlation, exporting history matters. This finding supports the basis premise of the hysteresis literature that there are substantial sunk costs involved in entering or exiting the export market.

Focusing on individual coefficients, we find that last year's exporting status $Y_{i,t-1}$ has a strong positive effect on the probability of exporting this year. But plants that last exported two or three years ago enjoy only small lingering effects from their previous investments in foreign-market access. The coefficients on $\tilde{Y}_{i,t-j}$ ($j = 2, 3$), which measure their respective discounts from the full sunk costs faced by a new exporter, are positive and decline with j as expected. However, neither parameter is significant, and we cannot reject the

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

[22] 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 0.795 percent. The differences, however, do not alter the general conclusion that transition rates, particularly for plants that do not export, are low and that 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.

[23] A more general framework would treat each plant as making simultaneous decisions to enter or exit production and to 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.

[24] In the estimates reported here we have restricted the serial correlation parameters ρ and ρ^p to be equal. We estimated a number of models that allowed a more general autoregressive process, but these were sufficiently overparameterized that none of the estimated autoregressive parameters was ever significantly different from zero. We also estimated models that allowed sunk costs to vary over time and with observable plant characteristics by including interactions between lagged participation and year dummies, plant size, business type, and industry. None of these additional interactions was statistically significant, and we do not report them here.

TABLE 3—DYNAMIC PROBIT MODEL OF EXPORT PARTICIPATION (STANDARD ERRORS IN PARENTHESES)

	Model 1	Model 2		Model 3	
Explanatory variable	(i) MSM	(ii) MSM	(iii) ML	(iv) MSM	(v) ML
Intercept	−7.105*	−7.058*	−7.039*	−6.856*	−7.033*
	(1.222)	(1.215)	(1.021)	(1.149)	(0.988)
Y_{t-1}	1.036*	0.971	1.140*	0.702*	0.885*
	(0.326)	(0.261)	(0.211)	(−0.154)	(0.135)
$\tilde{Y}_{t-2}$	0.326	0.331	0.401*	—	—
	(0.190)	(0.181)	(−0.145)		
$\tilde{Y}_{t-3}$	0.069	0.068	0.130	—	—
	(0.182)	(0.176)	(−0.164)		
1985 dummy	−0.156	−0.160	−0.168	−0.140	−0.154
	(0.133)	(0.109)	(0.106)	(0.100)	(0.097)
1986 dummy	−0.013	−0.022	−0.026	−0.017	−0.021
	(0.111)	(0.106)	(0.108)	(0.093)	(0.098)
1987 dummy	−0.309*	−0.312*	−0.340*	−0.286*	−0.318*
	(0.109)	(0.107)	(0.112)	(0.098)	(0.103)
1988 dummy	−0.148	−0.161	−0.187	−0.155	−0.178
	(0.119)	(0.115)	(0.114)	(0.102)	(0.104)
1989 dummy	−0.313*	−0.322*	−0.355*	−0.305*	−0.343*
	(0.111)	(0.110)	(0.118)	(0.098)	(0.109)
ln(Wage$_{t-1}$)	0.142	0.136	0.174	0.115	0.173
	(0.127)	(0.126)	(0.107)	(0.116)	(0.101)
ln(Export price$_{t-1}$)	−0.029	−0.027	−0.065	−0.025	−0.059
	(0.055)	(0.055)	(0.046)	(0.052)	(0.047)
ln(K_{t-1})	0.207*	0.211*	0.222*	0.207*	0.221*
	(0.032)	(0.032)	(0.031)	(0.031)	(0.033)
ln(Age$_{t-1}$)	0.471*	0.471*	0.349*	0.506*	0.396*
	(0.126)	(0.126)	(0.096)	(0.123)	(0.089)
Corporation	0.383*	0.386*	0.271*	0.450*	0.308*
	(0.156)	(0.152)	(0.115)	(0.148)	(0.111)
Textiles industry dummy	0.817*	0.815*	0.681*	0.839*	0.698*
	(0.159)	(0.159)	(0.135)	(0.158)	(0.158)
Paper industry dummy	0.310	0.305	0.165	0.319*	0.163
	(0.184)	(0.183)	(0.147)	(0.191)	(0.145)
Chemical industry dummy	0.762*	0.760*	0.640*	0.811*	0.689*
	(0.203)	(0.199)	(0.134)	(0.200)	(0.130)
Cali/Medellín	0.112	0.119	−0.017	0.119	−0.052
	(0.131)	(0.133)	(0.117)	(0.135)	(0.129)
Other region	0.479*	0.487*	0.453*	0.504*	0.471*
	(0.134)	(0.135)	(0.104)	(0.137)	(0.098)
Var(α)	0.668*	0.687*	0.620*	0.764*	0.688*
	(0.119)	(0.096)	(0.078)	(0.061)	(0.051)
Corr(α,α^P)	0.898*	0.894*	0.899*	0.906*	0.901*
	(0.039)	(0.038)	(0.017)	(0.032)	(0.017)
ρ	−0.019	—	—	—	—
	(0.028)				
Log likelihood:	−854.8[a]	−854.1[a]	−837.7	−857.8[a]	−842.700

* Statistically significant at the 5-percent level.
[a] Simulated with 1,000 draws of the errors.

hypothesis that both coefficients are jointly equal to zero. The $\chi^2_{[2]}$ statistic for the joint test is 3.18. Hence our choice of a three-year lag structure appears to be more than adequate to capture all of the relevant history.

B. *Expected Profits from Exporting*

The next group of coefficients in Table 3 summarizes the influence of year effects and plant characteristics on the expected profitability of exporting, net of sunk entry costs for a plant with no prior foreign market experience ($\pi_{it}^{*} - F_i^0$). The time dummies indicate there is variation over time, but only the 1987 and 1989 coefficients are significantly different from zero.[25] Net of entry costs, the expected future profits from exporting were highest in 1984 and 1986. This pattern corresponds to the high entry rates documented in Table 2 and may reflect producer anticipation of the coming years of favorable exchange rates (Table 1). But the incentives to begin exporting appear to soon dissipate, and by 1989 expected net profits reach their in-sample low. The fact that the estimated temporal shifts in expected future exporting profits do not closely track the exchange rate suggests that sunk costs and expectations play a critical role in shaping behavior.

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).[26] 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 curve.[27] Even if the annual payoff from exporting were the same for young and old plants, the young ones would perceive smaller returns to breaking into the market because they are less likely to survive.

Location matters as well, presumably because of transport costs. Bogotá, the base category in the model, is landlocked in the Andes mountain range, and plants there are among the least likely to produce for foreign markets. Cali and Medellín are also inland, but in regions that are less mountainous and closer to the coast. Nonetheless, we estimate that these cities are as unlikely to serve as a base for exporters as Bogotá. Perhaps their locational advantage is offset by their lack of Bogotá's agglomeration economies. Finally, the port cities of Cartagena and Barranquilla are the 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 do not 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, our sector-level export price indexes may do a poor job of summarizing foreign demand for the products of the individual producers in our sample. Similarly, measurement problems arise with our wage variable. It is constructed as the total cost of labor divided by the number of employees, so cross-plant variation in unit labor costs partly reflects variation in labor quality.

C. *Unobserved Plant Heterogeneity and Noise*

The final sources of variation in export status are unobserved error components:

[25] The hypothesis that the time dummies are jointly equal to zero is rejected with a Wald test. The $\chi^2_{[5]}$ test statistic is 14.59, and has a p-value of 0.012.

[26] Combined with our finding that sunk costs do *not* rise with size (not reported in Table 3), it is consistent with the well-known positive correlation between plant size and export participation. See Richard E. Caves (1989) and R. Albert Berry (1992) for a review of the evidence relating size and propensity to export.

[27] Roberts (1996) reports a decline in the probability of failure as a plant ages for Colombia, and Tybout (1996) reports a similar finding for Chile. Lili Liu and Tybout (1996) also find that failing plants in Colombia are systematically less productive than surviving plants. Both patterns have been found in data from the United States (see Timothy Dunne et al., 1989; Martin Neil Baily et al., 1992).

persistent plant heterogeneity, α_i, and transitory noise, ω_{it}. As shown in Table 3, $\mathrm{Var}(\alpha) = 0.668$, implying that two-thirds of the total unobserved variation is due to persistent heterogeneity, and this fraction is significantly different from 0 and 1. The correlation between the permanent component in the sample and presample years is very precisely estimated as $\mathrm{Corr}(\alpha, \alpha^{\mathrm{p}}) = 0.898$, indicating that our initial-conditions correction is critical.[28] However, once these sources of persistence in the error are controlled for, there is no evidence of additional serial correlation arising through correlated transitory shocks. The estimated serial correlation parameter ρ equals -0.019 with a standard error of 0.028.

D. *Alternative Model Specifications*

Given that serial correlation in transitory shocks is almost exactly zero, MSM estimates of the model are virtually identical when we impose the constraint $\rho = 0$ in column (ii) of Table 3. This constrained model can also be estimated with ML, yielding some efficiency gain, as shown in column (iii). Relative to our MSM estimates, the ML estimates attribute a slightly larger role to the plant's exporting history and a slightly smaller role to the plant characteristics age, ownership type, and industry. The standard errors of the ML estimates are slightly smaller than the MSM standard errors, as expected.

Together these changes result in a statistically significant coefficient on $Y_{i,t-2}$ when using the ML estimate. Thus, the ML estimates indicate that sunk costs are important and that costs incurred do not fully depreciate upon exit from the export market. Plants that have been out of the export market for one year do not incur reentry costs that are as substantial as those incurred by first-time entrants. However, the insignificant coefficient on $\tilde{Y}_{i,t-3}$ indicates that once the plant has been out of the market for two years the reentry costs are not significantly different from the costs of a first-time exporter.

The final two columns of Table 3 report MSM and ML estimates of a model which eliminates the more distant values of past participation. The main change resulting from this restriction is that the unobserved heterogeneity parameter, $\mathrm{Var}(\alpha)$, increases in magnitude as do the coefficients on a few of the plant characteristics. On the basis of a likelihood-ratio test using the ML estimates in columns (iii) and (v), we reject the restriction that the second and third lags have no explanatory power. The $\chi^2_{[2]}$ test statistic is 10.0 which has a p-value of 0.007. Thus, the ML estimates provide evidence of significant entry costs that increase with the length of time the plant has been out of the export market. Based on the results in Table 3 we have chosen the ML estimates of model 2 as the preferred set of results and will use these estimates in the analysis that follows.

E. *Goodness of Fit*

To assess the overall fit of this model we compare the actual and predicted patterns of export-market participation. Overall, given the six-year period 1984–1989 there are $2^6 = 64$ possible patterns or trajectories of export-market participation for an individual plant. Many of these patterns never occur or occur very infrequently among the 650 plants in our data set. To simplify comparison, we therefore aggregate the 64 trajectories into six categories based on whether we observe the plant in or out of the export market in 1984 and whether there were zero, one, or more than one transitions in or out over the remaining five years. Actual frequencies for each of the six categories are summarized in Table 4, along with predicted frequencies based on 200 simulations of the model for each plant.[29] Note that the actual and predicted frequencies are quite close. More formally, using Andrews's (1988) omnibus specification test to compare actual and predicted frequencies, we obtain a $\chi^2_{[5]}$

[28] In the presample years, the fraction of variation due to unobserved plant heterogeneity, $\mathrm{Var}(\alpha^{\mathrm{p}})$, equals 0.984. This occurs because the lagged participation variables are not used to predict Y_{it} during 1981–1983, and their effect is shifted to the disturbance.

[29] Actual and predicted frequencies for the complete set of 64 trajectories are reported in the Appendix.

TABLE 4—OBSERVED VERSUS PREDICTED FREQUENCIES OF Y_{IT} TRAJECTORIES (BASED ON ML ESTIMATES OF MODEL 2 IN TABLE 3)

Trajectory type	Observed frequencies	Predicted frequencies
Always a nonexporter	0.763	0.737
Begin as a nonexporter, switch once	0.045	0.044
Begin as a nonexporter, switch at least twice	0.029	0.033
Always an exporter	0.109	0.116
Begin as an exporter, switch once	0.022	0.037
Begin as an exporter, switch at least twice	0.032	0.034

statistic of 5.96, which has a p-value of 0.31.[30] Hence our functional form, error structure, and exogeneity specification appear to be appropriate, and the model appears to do a good job of predicting patterns of export-market participation.

F. *Sunk Costs, Heterogeneity, and Export Probabilities*

In order to assess the relative importance of observable plant characteristics, unobserved heterogeneity, and past participation on current export-market participation, we generate predicted probabilities of exporting for different types of plants. These probabilities, which are based on the ML estimates of model 2 in Table 3, are summarized in Table 5. The three panels going across the table allow the observable determinants of export profitability (age, capital stock, industry, etc.) to vary. Plants at the 25th, 50th, and 75th percentile of $\boldsymbol{\beta Z}_{it}$ are compared. Within each panel, plants are distinguished by whether they never exported ($Y_{t-1} = \tilde{Y}_{i,t-2} = \tilde{Y}_{i,t-3} = 0$), last exported three years earlier ($Y_{t-1} = \tilde{Y}_{i,t-2} = 0$; $\tilde{Y}_{i,t-3} = 1$), last exported two years ago ($Y_{t-1} = \tilde{Y}_{i,t-3} = 0$; $\tilde{Y}_{i,t-2} = 1$), or exported last year ($Y_{t-1} = 1$; $\tilde{Y}_{i,t-2} = \tilde{Y}_{i,t-3} = 0$). 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 standard deviations from zero.

Export history matters for plants that have either above-average $\boldsymbol{\beta Z}_{it}$ or above-average α_i values. For example, a plant with observable characteristics that place it at the 50th percentile of $\boldsymbol{\beta Z}_{it}$ and an unobserved plant effect equal to 1 will have a 0.052 probability of exporting if it was not an exporter in the past, but a 0.588 probability if it exported in the previous year. Among this group, participation in the previous year can increase the probability of exporting in the current years by as much as 0.63. On the other hand, for plants with below-average values of $\boldsymbol{\beta Z}_{it}$ or α_i, the expected profits from exporting 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 5 also demonstrates how quickly experience depreciates. Among plants that fall in the last two rows of the table, the difference in export probabilities between a plant that exported last year and an otherwise comparable plant that last exported two years earlier ranges from 0.11 to 0.45. The difference between plants that last exported two years ago and plants that last exported three years ago is substantially smaller, ranging from 0.03 to 0.17, and is based on a parameter estimate that is not statistically significant.

Finally, inferences about industry effects and macro conditions can also be drawn from Table 5. Specifically, an increase in α of one standard deviation corresponds to a $(0.668)^{0.5} = 0.82$ shift in expected profits from exporting, so anything that shifts expected profits from exporting by 0.82 corresponds to a one-row downward movement in Table 5. Hence, from the estimated industry dummies in Table 3, plants in the textile or chemicals industry that are otherwise comparable to plants in the food industry are approximately described by rows with plant effects one-standard-deviation larger. Notice that this type of downward movement in the table generally increases sensitivity to history because it concentrates

[30] The details of the test statistic are provided in the Appendix. Since each of the predicted frequencies is based on the same vector of coefficients, they are not independent, and a simple chi-square contingency-table test is inappropriate. Nonetheless, the values of this much simpler test statistic (not reported) are very close to the Andrews's test statistics in our applications.

TABLE 5—PREDICTED PROBABILITY OF EXPORTING
(BASED ON ML ESTIMATES OF MODEL 2 IN TABLE 3)

Plant effect (α)[a]	25th percentile of $\boldsymbol{\beta Z}_{it}$				50th percentile of $\boldsymbol{\beta Z}_{it}$				75th percentile of $\boldsymbol{\beta Z}_{it}$			
$(Y_{t-1}, \tilde{Y}_{t-2}, \tilde{Y}_{t-3}) =$	(0,0,0)	(0,0,1)	(0,1,0)	(1,0,0)	(0,0,0)	(0,0,1)	(0,1,0)	(1,0,0)	(0,0,0)	(0,0,1)	(0,1,0)	(1,0,0)
−2	0.000	0.000	0.000	0.003	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
−1	0.000	0.000	0.000	0.001	0.000	0.000	0.000	0.010	0.000	0.001	0.004	0.073
0	0.000	0.000	0.001	0.037	0.002	0.004	0.012	0.146	0.022	0.035	0.085	0.431
1	0.009	0.016	0.044	0.306	0.052	0.079	0.165	0.588	0.228	0.296	0.462	0.865
2	0.141	0.193	0.335	0.780	0.363	0.445	0.618	0.933	0.702	0.771	0.881	0.991

Note: Each table entry is the predicted probability of exporting for a given combination of export history $(Y_{t-1}, \tilde{Y}_{t-2}, \tilde{Y}_{t-3})$, plant effect ($\alpha$), and export profitability (percentile of $\boldsymbol{\beta Z}_{it}$).
[a] Measured as standard deviations from the mean.

more plants near their action thresholds. For example, an industry in which half of the plants are already exporting will be more responsive to a one-unit rightward shift in the expected profit distribution than industries in which only 5 percent of the plants are exporting, since a larger fraction of firms in the former industry will find their expected profits pushed above sunk entry costs. One implication of this phenomenon is that changes in the exchange rate and other measures that generally promote exports will not scale up participation proportionately in all industries. Responses are more likely to be concentrated in a few sectors.

Overall, the results reported in Tables 3 and 5 reveal that the export-participation decision is affected by time-period or macroeconomic 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 last factor has a particularly substantial effect on the probability that a plant exports, and this is consistent with the plant's facing significant entry costs in the export market.

V. Conclusions

To explain the unpredictable responses of trade flows to exchange rate movements, recent theoretical models have stressed that producers face sunk entry costs when breaking into foreign markets. These imply, for example, that devaluations that induce entry into the export market may permanently increase the flow of exports, even if the currency subsequently appreciates. On the other hand, conditions that appear favorable to exporting 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. Finally, the combination of sunk costs and uncertainty about future market conditions can create an option value to waiting.

While the distinguishing feature of these models is their focus on market entry and exit, empirical tests to date have relied on aggregate or sectoral data on trade flows and prices. Thus the sunk-cost explanation for hysteresis in trade flows has remained plausible but untested. In this paper we have developed an econometric model of plants' decisions to diversify into new markets and used it to analyze plant-level exporting decisions for consistency with the theory. Using panel data on a large group of Colombian manufacturing plants we reject the null hypothesis that entry costs are unimportant, and this implies that sunk-cost hysteresis models are empirically relevant.

Our empirical results also reveal that exporting experience depreciates once plants cease servicing foreign markets. After a two-year absence the reentry costs are not significantly different from 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. For plants with "average" observable characteristics and no prior exporting experience, variation in *un*observed sources of difference in profitability can lead to as much as a 36-percentage-point difference in the probability of exporting. For other plants the effect can be much larger.

This combination of plant heterogeneity and sunk costs implies that the response of aggregate or sectoral exports to changes in policy or the macroeconomic 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 existing 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 macroeconomic 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.

APPENDIX: SPECIFICATION TESTS

Andrews's (1988) specification test compares realized $(Y_{it}, \mathbf{Z}_{it})$ trajectories in our sample to expected trajectories based on the estimated model. To construct his chi-square statistic, we first partition the possible $(Y_{it}, \mathbf{Z}_{it})$ 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 (Table A1). To avoid cells that are nearly empty we distinguish six types of trajectories: $Y_{it} = 0 \quad \forall\, t$; $Y_{it} = 0$ initially, but switches *once* during the sample period; $Y_{it} = 0$ initially and switches *at least twice* during the sample period; $Y_{it} = 1 \quad \forall\, t$; $Y_{it} = 1$ 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 $\mathbf{\Psi}(Y_{i,J+1}, Y_{i,J+2}, \ldots, Y_{i,J+T})$ map Y_{it} sequences into these six cells, the observed frequencies of the different trajectories in our sample is the vector

$$\mathbf{P}_n(\mathbf{\Psi}) = \frac{1}{n} \sum_{i=1}^{n} \mathbf{\Psi}(Y_{i,J+1}, Y_{i,J+2}, \ldots, Y_{iT}).$$

Next, to generate model-based expected values for each of these cells, we use estimated parameter values from the ML estimates of model 2 in Table 3 in conjunction with the observed $\mathbf{Z}_{it}$ trajectories and random draws on α_i, 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 ith plant, we get the probabilities that it will fall in each cell under the null hypothesis that our specification is correct: $\mathbf{Q}(\mathbf{\Psi}, \mathbf{Z}_{iJ}, \mathbf{Z}_{i,J+1}, \ldots, \mathbf{Z}_{iT} \mid \mu, \boldsymbol{\beta}, \boldsymbol{\gamma}, \sigma_\alpha, \rho)$. Finally, averaging these probabilities over all plants, we obtain the expected sample-wide frequencies of each cell in the partition:

$$\mathbf{Q}_n(\mathbf{\Psi}) = \frac{1}{n} \sum_{i=1}^{n} \mathbf{Q}(\mathbf{\Psi}, \mathbf{Z}_{iJ}, \mathbf{Z}_{i,J+1}, \ldots, \mathbf{Z}_{iT} \mid \mu, \boldsymbol{\beta}, \boldsymbol{\gamma}, \sigma_\alpha, \rho).$$

TABLE A1—PREDICTED AND ACTUAL FREQUENCIES OF Y_{it} TRAJECTORIES

Predicted frequency	Actual frequency	Export status[a]					
		1984	1985	1986	1987	1988	1989
0.116	0.109	1	1	1	1	1	1
0.007	0.003	1	1	1	1	1	0
0.003	0.002	1	1	1	1	0	1
0.004	0.003	1	1	1	1	0	0
0.007	0.005	1	1	1	0	1	1
0.001	0	1	1	1	0	1	0
0.002	0	1	1	1	0	0	1
0.007	0.005	1	1	1	0	0	0
0.002	0.003	1	1	0	1	1	1
0.000	0	1	1	0	1	1	0
0.000	0	1	1	0	1	0	1
0.000	0.002	1	1	0	1	0	0
0.001	0.003	1	1	0	0	1	1
0.000	0	1	1	0	0	1	0
0.001	0.002	1	1	0	0	0	1
0.006	0.002	1	1	0	0	0	0
0.005	0.006	1	0	1	1	1	1
0.001	0	1	0	1	1	1	0
0.000	0.002	1	0	1	1	0	1
0.001	0	1	0	1	1	0	0
0.001	0	1	0	1	0	1	1
0.000	0	1	0	1	0	1	0
0.000	0	1	0	1	0	0	1
0.001	0.006	1	0	1	0	0	0
0.001	0	1	0	0	1	1	1
0.000	0	1	0	0	1	1	0
0.000	0	1	0	0	1	0	1
0.000	0	1	0	0	1	0	0
0.002	0.002	1	0	0	0	1	1
0.000	0.002	1	0	0	0	1	0
0.001	0	1	0	0	0	0	1
0.012	0.009	1	0	0	0	0	0
0.008	0.014	0	1	1	1	1	1
0.002	0	0	1	1	1	1	0
0.000	0	0	1	1	1	0	1
0.001	0	0	1	1	1	0	0
0.001	0	0	1	1	0	1	1
0.000	0	0	1	1	0	1	0
0.001	0	0	1	1	0	0	1
0.002	0.002	0	1	1	0	0	0
0.000	0	0	1	0	1	1	1
0.000	0	0	1	0	1	1	0
0.000	0	0	1	0	1	0	1
0.000	0	0	1	0	1	0	0
0.000	0.002	0	1	0	0	1	1
0.000	0.002	0	1	0	0	1	0
0.000	0	0	1	0	0	0	1
0.002	0.003	0	1	0	0	0	0
0.009	0.006	0	0	1	1	1	1
0.002	0	0	0	1	1	1	0
0.001	0	0	0	1	1	0	1
0.002	0.002	0	0	1	1	0	0
0.001	0	0	0	1	0	1	1
0.001	0	0	0	1	0	1	0
0.001	0	0	0	1	0	0	1
0.006	0.012	0	0	1	0	0	0
0.005	0.008	0	0	0	1	1	1
0.002	0	0	0	0	1	1	0
0.000	0	0	0	0	1	0	1
0.002	0.003	0	0	0	1	0	0
0.011	0.011	0	0	0	0	1	1
0.005	0.005	0	0	0	0	1	0
0.011	0.006	0	0	0	0	0	1
0.737	0.763	0	0	0	0	0	0

[a] Export-market participation is denoted by 1's; nonparticipation is denoted by 0's.

The test statistic is calculated as a quadratic form in the difference between the two columns, $\chi^2_{[K-1]} = (\mathbf{P}_n - \mathbf{Q}_n)'\hat{\mathbf{W}}(\mathbf{P}_n - \mathbf{Q}_n)$; where the weighting matrix $\hat{\mathbf{W}}$ is given by equation (15) in Andrews's (1988) Appendix.

Using the ML estimates of model 2 in Table 3 we also generate the predicted frequency of each of the 64 possible trajectories. These are reported in the first column of Table A1, and the actual frequencies are reported in the second column. Andrews's (1988) test statistic can be used to compare the full set of 64 predicted and actual trajectories, although the fact that many trajectories never occur makes this exercise suspect. Doing so, we find a $\chi^2_{[63]}$ value of 83.9 which has a p-value of 0.04.

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