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IMPORT COMPETITION AND THE STOCK MARKET RETURN TO CAPITAL

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ABSTRACT

We measure the responsiveness of returns to capital invested in six U.S. industries to shocks to the prices of competing import goods. Recognizing that most capital services are not traded on spot rental markets, we treat the intersectoral mobility of capital as the outgrowth of investment behavior. Then the return to capital is realized as an asset return to equity holders. We model expected returns by CAPM, and relate "excess" returns in a period to unanticipated shocks to the variables that affect current and future profits. We find that positive shocks to import prices cause higher than normal stock market returns in all six industries. The magnitudes of the responses are consistent with the hypothesis that capital is highly sector specific in five of these industries.

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I. Introduction

The Stolper-Samuelson derivatives hold a central place in the theory of international trade. These parameters, measuring the sensitivity of domestic factor prices to changes in the output prices of internationally traded goods, reveal the distributional implications of terms of trade changes, and suggest the political economic motivations for trade and industrial policies.

At least since Jones (1971), Mayer (1974), and Mussa (1974), trade economists have been well aware of the importance of factor specificity in determining the effect of commodity price changes on factor rewards. When factors are mobile, as in the analysis of the Heckscher-Ohlin model by Stolper and Samuelson (1941) and Jones (1965), factor returns are governed by conditions characterizing the full general equilibrium of the economy, and individual returns may respond little or even positively to adverse shocks to the particular sectors in which the factors are employed. By contrast, when a factor is specific to a particular activity, its fate is tied closely to that of its industry of employment.

Surprisingly, there have been few attempts to measure the sensitivity of factor returns to international conditions, or to assess the intersectoral mobility of factors. Magee (1980) provides some indirect evidence on this from observed lobbying behavior of industry associations and trade unions. He reports that, for a cross-section of industries, capital and labor employed in an industry are much more likely to hold similar views on the desirability of freer trade than they are to hold opposing views. This finding, he notes, supports the hypothesis that factors are specific to their sector rather than intersectorally mobile. A small number of studies have attempted more direct measurement of the effects of import competition on factor returns, but these have been limited, as far as we know, to the investigation of wage responses.[1]

Our goal in this paper is to measure the responsiveness of the returns to capital invested in a number of U.S. industries to shocks to the prices of competing import goods, and to infer therefrom a sense of the intersectoral mobility of capital. Direct application of the familiar trade models would seem to suggest a procedure for accomplishing this. That is, we might think to regress the rental rate of capital on current and lagged values of the variables that theory tells us should affect factor returns, including among others the price of foreign products. Indeed, this is the procedure that Grossman (1987) followed in his study of wage responsiveness. Unfortunately, in attempting to implement this procedure, we immediately confront the fact that most capital services are not traded on spot rental markets, as is typically assumed in theoretical models. Rather, capital goods most often are purchased outright by firms and installed as fixed equipment, so that the return to capital is realized as an asset return to equity (and debt) holders. Our attention must focus, therefore, on the determination of asset returns on highly efficient, forward-looking, financial markets.

Our approach here bears some resemblance to a recent study by Pakes (1985) of the relationship between R&D expenditures, patent applications, and the stock market returns on firms' equities.[2] We treat capital installation using a simple theory of investment, with the degree of capital mobility captured by a parameter in the cost-of-adjustment function. Expected stock

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[1] See, for example, Grossman (1987), Abowd (1987) and Heywood (n.d.).

[2] See also Gardner (1986), who adopts a similar approach to study the vulnerability of firms in the scientific instruments industry to exchange rate fluctuations. More generally, our analysis bears some resemblance to the event-study methodology; see Schwert (1981).

returns are modelled by CAPM, with excess returns measured by the deviations of realized returns from the predictions of the CAPM equation. Finally, we relate excess returns to unanticipated shocks (innovations) in the variables that affect current and future profitability of the firm, including among these the domestic currency price of competing foreign products. This approach reflects our belief that only unanticipated changes in the extent of import competition should have measurable effects on realized equity returns; forward-looking traders already will have capitalized the implications of expected changes in profitability variables into the beginning-of-period values of the shares.

The remainder of the paper is organized as follows. In Section II we explore the theoretical relationship between import competition, asset values, and the intersectoral mobility of capital. In Section III we discuss the elaboration of the model necessary for empirical application. Data and estimation issues are treated in Section IV. Section V contains our primary findings and interpretation. Finally, we present some sensitivity analysis in Section VI and conclude in Section VII.

II. Import Competition and Stock-Market Returns: A Theoretical Framework

To explore the theoretical relationship between import competition, stock-market values, and the degree of intersectoral capital mobility, we adopt a simple, dynamic, competitive model of capital formation, output production, and industry equilibrium. We follow Mussa (1978) in treating intersectoral movements of capital as the outgrowth of investment decisions by firms. But our analysis is simpler than his, since a partial-equilibrium framework is sufficient for our purposes.

Consider then a competitive industry in which imports substitute imperfectly for home goods. The home country is assumed to be small, so that the time path of import prices, p_t^*, can be taken to be exogenous. Domestic output is produced with capital and a vector of intersectorally mobile factors according to a constant-returns-to-scale production function. All home firms have access to the same technology.

We take the capital stock of firm i, K_i, to be a state variable, alterable by the installation or removal of fixed machinery and equipment. For the purposes of this discussion, we ignore depreciation, so that $\dot{K}_i = I_i$, where I_i is investment by firm i. Investment (or disinvestment) involves two costs. First, the capital equipment must be purchased (or sold, in the case of disinvestment) at the fixed price of one dollar per unit of capital. Second, there is a convex installation (or dis-installation) cost that limits the extent of investment at any point in time. We assume, for simplicity, that these adjustment costs are symmetric for positive and negative investment, and that they take the particular form, $\gamma I_i^2/2K_i$. In this expression, γ is a shift parameter that raises or lowers uniformly both the total and marginal costs of investment. We take γ as our measure of the intersectoral mobility of the industry's capital, because it indicates the ease with which capital can be moved into or out of the sector.

We assume that the equities of the firms in the industry are traded on a perfect, efficient asset market. We assume as well (in this section, but not in our empirical work) that investors are risk neutral, and that there exists a risk-free asset paying a rate of return r_f. Since we are in an environment where the Modigliani-Miller theorem holds, the choice of investment financing by debt versus equity issue has no real effects. Also, with tax considerations absent from the model, a firm's dividend policy is irrelevant. So we can assume with no further loss of generality that the firm pays out its current cash flow as dividends to its stockholders. Then our assumptions imply that the value of the firm, V_i, is the present discounted value of the firm's future cash flow, where discounting is at rate r_f.

Cash flow is the difference between operating profits and investment costs. By the twin assumptions of constant returns to scale and perfect mobility of all factors other than capital, we can write operating profits as $K_{it}\pi(p_t)$, where $\pi(\cdot)$ represents the maximized value of instantaneous profits per unit of capital and the maximization is taken over the choice of variable inputs. Then the value of firm i at time t is

$$V_i(t) = \max \int_t^\infty \left\{ \pi(p_t)K_{it} - I_{it} - \frac{\gamma I_{it}^2}{2K_{it}} \right\} e^{-r_f t}\, dt \quad . \tag{1}$$

Instantaneous market clearing determines the path of domestic prices used by the firm in maximizing its value in (1). The condition for industry equilibrium can be written as $p_t = \phi(x_t, p_t^*)$, where $x_t = \Sigma_i K_{it}\pi'(p_t)$ is aggregate output by home firms, and $\phi_1 < 0$, $\phi_2 > 0$.

To find the optimal path for the capital stock of firm i, assuming perfect foresight about import prices, we form the current value Hamiltonian

$$H_i = \pi(p)K_i - I_i - \gamma I_i^2/2K_i - \lambda_i(\dot{K}_i - I_i) \ .$$

The first order condition for optimal investment implies

$$\dot{K}_i = I_i = (\lambda_i - 1)K_i/\gamma \ . \tag{2}$$

The co-state variable evolves according to

$$\dot{\lambda}_i = r_f \lambda_i - \pi(p) - \gamma I_i^2/2K_i^2 \ . \tag{3}$$

Notice from (2) that I_i/K_i depends only on λ_i, while (3) reveals that the evolution of λ_i depends only on I_i/K_i. It follows that I/K and λ will be identical across firms.

Let us now suppose that agents expect the import price to remain constant forever at some level, p*. The dynamic evolution of the state and co-state variables for this case can be shown in the familiar phase diagram of Figure 1. We draw the $\dot{\lambda} = 0$ schedule as downward-sloping to reflect the fact that an increase in the capital stock of firm i corresponds to an equiproportionate increase in the industry-wide capital stock (since all firms follow similar investment behavior), and therefore to an increase in output and a fall in p. As usual, there exists a unique path for the industry equilibrium that is stable and converges to the steady state. This saddlepath for the dynamic system is depicted in the figure.

What will happen if, having achieved a steady-state equilibrium corresponding to p^*, the industry is shocked by an unexpected abatement of import competition? That is, suppose that the import price jumps suddenly to $p^{*\prime}$, a change that is perceived to be permanent. This shock causes the $\dot{\lambda} = 0$ schedule to shift out, as shown in Figure 2. The system now is governed by a new set of differential equations, and begins to move to the new steady-state equilibrium.

Notice that the co-state variable jumps at the instant of the shock, but that it adjusts smoothly to its new steady-state value thereafter. By a theorem of Hayashi (1981, p.218), we can relate the co-state variable to the value of the firm. Hayashi proves that with perfect competition, constant returns to scale, and an adjustment-cost function that is linearly homogeneous in I and K, Tobin's marginal q is equal to average q. Marginal q in our model is just equal to λ, because the purchase price of capital is taken to be unity. Average q is V_i/K_i. Thus, the jump in λ at the time of the shock corresponds exactly to the jump in the value of the firm per unit of installed capital. It follows that equity holders earn an abnormal return (positive, in this case) at the moment that the "news" about import competition is learned. Thereafter, the higher than normal dividends that are realized while the capital stock is growing are offset by perfectly anticipated capital losses on the value of the shares. Total equity returns during the adjustment path are "normal", i.e., just equal to r_f.

Finally, we are ready to investigate the role of capital mobility in determining the sensitivity of asset returns to shocks to import prices. Let us compare the effect of similar shocks in two industries that differ only in the size of γ. Notice that the initial and final steady-state points in Figure 2 are independent of γ, but that the path between them is not. We show formally in Appendix A that, at least in the neighborhood of the new steady-state equilibrium, the saddlepath for a firm in the industry with higher installation costs must be steeper than that for a similarly sized firm in the industry with more mobile capital. Then, if the change in p^* is not "too" large, the value of λ at the moment after the shock must be larger for the industry with the higher value of γ. Since the initial capital stock is the same in both cases, it follows from the previously cited theorem of Hayashi

(1981) that the jump in stock prices is larger in the industry with less mobile capital. Specificity makes capital more vulnerable to shocks to the profitability of the industry.

We summarize the analysis of this section in terms of its implications for the time-series properties of the returns to shares of firms in a particular industry. In any time period, the expected return to these equities equals the expected return on the market portfolio, which, under the assumption of risk neutrality, also equals the risk-free rate. That is,

$$E\, r_{it} = E\, r_{mt} = r_f , \tag{4}$$

where r_{mt} is the market return. Realized returns for shares of firm i may differ from expected returns due to the influence of <u>unanticipated</u> shocks to variables that affect current and future profitability of that firm. Letting u_{it} reflect the total effect of all such shocks, we have

$$r_{it} = E\, r_{it} + u_{it} . \tag{5}$$

Similarly, for the market portfolio,

$$r_{mt} = E\, r_{mt} + v_{mt} . \tag{6}$$

Combining (5) and (6) gives us the relationship between the realized returns,

$$r_{it} = r_{mt} + u_{it} - v_{mt} . \tag{7}$$

Finally, we would expect the coefficient on any given component of u_{it} (for example, a permanent shock to import prices of a given magnitude) in (7) to be larger, the less mobile is the capital used in production in the industry.

III. Elaboration of the Model for Empirical Application

Under the assumption of risk neutrality, arbitrage ensures the simple relationship between expected returns on different assets given by (4). But a large body of research in financial economics rejects the hypothesis of risk neutrality as applied to asset markets. The risk characteristics of different assets are known to be important determinants of their expected returns. Before turning to the data, therefore, it is imperative that we extend our model to allow for uncertainty and risk aversion on the part of investors.

Certainly the most satisfactory way for us to incorporate risk into our model would be to introduce all the primitive sources of uncertainty (import prices, factor prices, technology and demand), and then to derive investment behavior and asset-pricing formulas from the multi-period utility maximization of consumer-investors. Unfortunately, multi-period, general-equilibrium asset pricing models that take as their starting point the stochastic processes of shocks to taste and technology are just now being developed by finance theorists (Gibbons, 1987, p.37). Those models that have been analyzed (e.g., Cox, Ingersoll and Ross, 1985) typically assume a one-good economy and suppress the role of factor markets. A suitable extension of such models to incorporate many goods and several factors, some of which are imperfectly mobile, would most likely yield a set of equations far too complex for empirical implementation, and in any event is beyond the scope of this paper.

Instead we choose a simpler, albeit somewhat more _ad hoc_, approach.[3] We assume that the relationship between the expected returns on different assets is as predicted by the _capital asset pricing model_ (CAPM):

$$E\, r_{it} - r_f = \beta_i (E\, r_{mt} - r_f) \tag{8}$$

In our sensitivity analysis of Section VI, we do allow for the possibility that the β_i in (8) vary over time, as might occur in response to movements in the state variables of the industry; however, even there, we approximate the movements of β_i by a quadratic time trend, rather than imposing the restrictions on the relationship between β_i and the other endogenous variables that a rigorous theoretical derivation would imply.

Combining (5), (6) and (8), we find the relationship between the realized returns that is implied by CAPM:

$$r_{it} = (1-\beta_i) r_f + \beta_i r_{mt} + u_{it} - \beta_i v_{mt} \tag{9}$$

Our next task is to specify the components of u_{it} and v_{mt}. Recall that u_{it} represents the combined impact on the realized return of news that is acquired during the period about variables that will affect current and future profits of the firm. We adopt a reduced-form, partial-equilibrium approach similar to that followed by Grossman (1987). We write reduced-form profits as a log-linear combination of variables that are exogenous to the industry, namely

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[3] Our approach is identical to that adopted in much of the event-study literature. See, for example, Rose (1985) and Hartigan, Perry and Kamma (1986). Gardner (1986), in his study of the effects of exchange-rate fluctuations, also adopts CAPM as a starting point.

economy-wide factor prices (wages and energy prices) and exogenous demand variables (aggregate income, prices of "other" goods, and the price of competing imports).[4] Then u_{it} comprises a linear combination of the unanticipated components of the realizations of these variables during period t, as well as the changes in beliefs about the future values of these variables that occur due to the updating of expectations during period t.

To identify the innovations in the variables that enter the reduced-form profit function, we posit the form of the stochastic process that each relevant variable follows, and then estimate the parameters from time-series data. Consider first the evolution of the domestic currency price of the import good. Competitive pricing behavior implies $p_t^* = e_t c_t^*(f_t^*)$, where e_t is the exchange rate at time t, c_t^* is the foreign-currency cost function for the import good prevailing at t, and f_t^* is a vector of factor prices. In principle, the exchange rate, the parameters of the production function, and the various factor prices might all follow different stochastic processes. Of course, we do not observe the technology parameters, nor is it practical for us (or the investor) to collect and incorporate data on all foreign factor prices in forming an expectation about p^*. Instead, we suppose that import prices contain a trend component (due, for example, to improvements in technology), and that they are influenced by own lagged values, and by lagged values of the exchange rate and foreign wages (w_t^*). More specifically, we assume that a detrended series for the log of p^* follows a multivariate autoregressive process given by:

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[4] The exogeneity of these variables relies on the assumption that the industry under consideration is small in relation to the U.S. economy, and that the United States is small in its import markets.

$$\tilde{p}^{*} = \tilde{p}_0^{*} + \Sigma_{j=1}^{4} \rho_{pj}\tilde{p}_{t-j}^{*} + \Sigma_{j=1}^{8} \rho_{wj}\tilde{w}_{t-j}^{*} + \Sigma_{j=1}^{8} \rho_{ej}\tilde{e}_{t-j} + PSNEWS_t \qquad (10)$$

where $PSNEWS_t$ is a white-noise error term, and the "tildes" denote detrended, deseasonalized logs of the respective variables.[5] We assume as well that the (detrended logs of) the foreign wage and the exchange rate follow univariate autoregressive processes, with four lagged own-values and residuals $WSNEWS_t$ and $ERNEWS_t$, respectively. Then, it is easy to show that $p_t^{*} - E_{t-1}p_t^{*} = PSNEWS_t$; and that, for all j, $E_t p_{t+j}^{*} - E_{t-1}p_{t+j}^{*}$ is a linear combination of $PSNEWS_t$, $WSNEWS_t$, and $ERNEWS_t$. Thus, a non-zero realization of any of the three variables, $PSNEWS_t$, $WSNEWS_t$, and $ERNEWS_t$, causes an updating of beliefs about current and future import prices and contributes to a deviation of the actual from the expected stock returns.

We generate forecasts about the other variables that affect current and future returns to industry capital (i.e., aggregate wages, energy prices, G.N.P., and aggregate producer prices) by assuming that the detrended logs of these variables follow a vector autoregressive (VAR) process that includes four lags of each of the variables and four lags of the money supply. The latter variable, while it does not affect profitability directly, is held to be useful as a predictor of the others.[6] We take the residuals from this vector autoregression to be components of u_{it} (see Table 1 for variable definitions), and write

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[5] All trends are assumed to take quadratic forms, so that the detrended variables are the residuals from regressions on time and time-squared. The time period in all our empirical analysis is one quarter. Section IV gives more details of the estimation. We consider the robustness of our results to alternative specifications of the process generating p^{*} in Section VI.

[6] The estimated coefficients from the VAR confirm our priors about the significance of the money supply as a leading indicator. This finding accords well with that of many researchers before us.

$$u_{it} = a_{1i}PSNEWS_t + a_{2i}WSNEWS_t + a_{3i}ERNEWS_t + a_{5i}PENEWS_t + a_{6i}WNEWS_t + a_{7i}GNPNEWS_t + a_{8i}PPINEWS_t + a_{9i}MSNEWS_t + \mu_{it} \qquad (11)$$

Here μ_{it} represents the combined effects of other information acquired by investors during period t that is unobservable to the econometrician. Note that each coefficient a_{ji}, for j=5,8, incorporates both direct and indirect impacts: for example, an innovation in the aggregate wage lowers profits in the current period, but also alters agents' predictions of the future values of the other variables that determine profits. Without detailed knowledge of both the structural parameters of the underlying model and the autoregressive parameters of the VAR process, we can have no strong priors about the signs of these reduced-form coefficients.

Our treatment of v_{mt} is similar. The same set of aggregate variables influences profitability throughout the economy. However, in place of innovations in p_t^* and the variables that help to forecast it, we include in v_{mt} innovations in the index of aggregate import prices, as a measure of shocks to economy-wide import competition. We model $AGGMNEWS_t$ as the residual from a fourth-order autoregression using detrended logarithms. Then,

$$v_{mt} = b_4AGGMNEWS_t + b_5PENEWS_t + b_6WNEWS_t + b_7GNPNEWS_t + b_8PPINEWS_t + b_9MSNEWS_t + \nu_{mt} \qquad (12)$$

Substituting (11) and (12) into (9), we have finally a relationship between the realized return on a particular stock, the market return, and innovations in the variables that either affect profits directly, or that influence the time-series evolution of variables that do:

$$
\begin{aligned}
r_{it} = \Gamma_0 &+ \beta_i r_{mt} + \Gamma_{1i}\mathrm{PSNEWS}_t + \Gamma_{2i}\mathrm{WSNEWS}_t + \Gamma_{3i}\mathrm{ERNEWS}_t + \Gamma_{4i}\mathrm{AGGMNEWS}_t \\
&+ \Gamma_{5i}\mathrm{PENEWS}_t + \Gamma_{6i}\mathrm{WNEWS}_t + \Gamma_{7i}\mathrm{GNPNEWS}_t + \Gamma_{8i}\mathrm{PPINEWS}_t \\
&+ \Gamma_{9i}\mathrm{MSNEWS}_t + \mu_{it} - \beta_i \nu_{mt} \ ; \qquad (13)
\end{aligned}
$$

where $\Gamma_0 = (1-\beta_i)r_f$ and $\Gamma_{ji} = a_{ji} - \beta_i b_j$ for $j = 5,9$.

Our main interest concerns the size of Γ_{1i}. This coefficient represents the closest empirically identifiable analog to the Stolper-Samuelson derivative of the simple, static models. A large positive value of Γ_{1i} would indicate, for example, that a lower-than-expected import price in the current period can cause substantial capital losses for the owners of firms with capital invested in the industry. Shareholders might well be expected to complain vociferously of the ill effects of import competition under such circumstances.

IV. Data and Estimation Issues

A. Data Sources and Methods

We sought to measure the sensitivity of stock market returns to foreign prices for as wide a range of U.S. import-competing industries as possible. Our criteria for selection of industries were as follows. First, we identified sectors for which a reasonably long time series of import prices was available. Import prices from survey data are published by the Bureau of Labor Statistics in U.S. Import and Export Price Indexes, but most series begin with quite recent observations. Longer series exist for a small number of categories, and these were the initial candidates for our study. We eliminated several categories that were primarily export industries. Among the remaining sectors, we chose all those for which no binding quantitative restrictions were in effect during the sample period. Since our method requires the assumption of a perfectly-elastic import supply, it would not be appropriate to apply it where trade is subject to quotas or export restrictions. This selection procedure left six industries. The industries are listed in Table 1; sample periods are shown at the bottom of Table 3.[7]

For each industry, we included in our sample all firms traded throughout the period on the New York Stock Exchange whose primary line of business, as reported in Ward's Business Directory of Largest U.S. Companies, coincides with the SIC category under consideration.[8] The number of firms for each product group is shown in Table 3. Monthly stock returns (dividends plus capital gains) were taken from the 1986 Master File of the University of Chicago's Center for Research in Security Prices (CRSP). The monthly returns were compounded to yield quarterly rates.[9]

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[7] The sample period for SIC category 331 was cut short in 1984:2, in recognition of the voluntary export restraints that took effect in October, 1984. For this category only, we extended the sample period back beyond the starting date of the BLS import price series, by using a carefully constructed index of unit values from Grossman (1986). The sample period for SIC 262 ends in 1985:4, as two corporate acquisitions that took place during 1986 would otherwise have eliminated several firms from our sample. With these exceptions only, our samples use all of the available data.

[8] The only exception to this rule concerns SIC 32, where we excluded firms active in the production of asbestos. Returns to these firms have been substantially affected by the evolution of the product liability lawsuits that took place during our sample period.

[9] Our procedure for compounding incorporates the implicit assumption that dividends are paid on the last day of the month, and that they are reinvested in the firm at the share price prevailing on that day. Pakes (1985) notes that a correction for a similar type of approximation was inconsequential for his findings.

The remainder of the data is from several sources.[10] Macroeconomic variables (non-agricultural wages, U.S. GNP, producer prices, energy prices and the M1 money supply) were taken from the Citibase Databank. The index of aggregate import prices (actually an index of unit values) is from the Survey of Current Business. Six separate indexes for foreign wages and for the U.S. exchange rate were constructed for use in predicting the various import prices. For each index, we geometrically weighted the wage series and the dollar exchange rate series for the major supplier countries of a particular commodity by the value shares of those countries in 1980 U.S. imports. End-of-period exchange rates are those reported in the I.M.F.'s International Financial Statistics. Wage data are from O.E.C.D., Main Economic Indicators.

B. Construction of the "News" Variables

We turn now to the construction of the news variables. The shortage of observations on import prices argues for the use of as parsimonious a specification of the process generating this variable as possible; yet the central role that this variable plays in our study dictates that we strive to minimize any measurement error in this series. To achieve these goals, we adopted a nested hypothesis-testing procedure. After quadratically detrending all the variables, we estimated the multivariate autoregressions (equation (9)) for each SIC category. We then tested separately for the joint significance of the second through fourth lags on $\tilde{p}^*$, the fifth through eighth lags on $\tilde{w}^*$, and the fifth through eighth lags on $\tilde{e}$. If we could not reject

[10] A detailed data appendix describing sources and methods of variable construction is available from the authors upon request. We will also make available to interested parties those portions of our data set not subject to copyright restrictions. Please include several 5.25-inch double-density, floppy disks with any request.

the hypothesis that each of these sets of lags was different from zero at the 90 percent confidence level, then we re-estimated the autoregression without the lags identified as insignificant. We repeated this procedure, testing and excluding as appropriate the second through fourth lags on $\tilde{w}^*$ and $\tilde{e}$, and finally the remaining (first) lag on $\tilde{w}^*$ and $\tilde{e}$. The results of all these F-tests, and the final lag structure adopted for each of the SIC categories, are shown in Table 2. The residuals from these final-specification autoregressions constitute our PSNEWS series for the various industries.

We initially specified the processes for the exchange rate, the foreign wage rate and the aggregate import price series as fourth-order autoregressions of the quadratically detrended series. We found however that the second through fourth-order terms often were not significant, either singly, or jointly. So, by nested-hypothesis testing, we pared down the specifications of these autoregressions until the coefficient on the last lag was significantly different from zero at the 80% confidence level. In each case, we checked that the resulting residuals showed no evidence of serial correlation. The sample period for the construction of AGGMNEWS was 1974:2 to 1986:4. Those for the various WSNEWS and ERNEWS variables were the same as for the corresponding PSNEWS variables, as reported in Table 2.

Finally, we estimated the fourth-order VAR using the five detrended aggregative variables. We used quarterly data from 1959:1 through 1986:3 for this estimation. No one of these variables was consistently insignificant across all regressions, nor were the later lags always insignificant. Consequently, we stayed with our initial specification in this case. The remaining news variables were created as the residuals to these VARs.

C. Estimation of the Coefficients of the Reduced-Form Return Equations

Estimation of the coefficients of the reduced-form return equations raises several econometric issues, among them a classic errors-in-variables problem, a heteroscedastic error term resulting from the pooling of observations for different firms in the same industry, and a potential bias in the estimates of the standard errors caused by the use of estimated residuals as both independent and dependent variables. We discuss each of these issues in turn.

Ordinary least squares applied to equation (13) would yield inconsistent estimates of most, if not all, of the parameters. The reason is the familiar errors-in-variables problem. The CAPM model specifies a relationship between expected returns, whereas the econometrician observes realized returns. The linear dependence of r_{mt} on v_{mt} is clear from the definition of the latter variable in (6). Then, ν_{mt}, which is the unobservable component of v_{mt} in (12), must also be correlated r_{mt}. Thus, the error term in (13), which includes ν_{mt}, must be correlated with one of the right-hand-side variables, unless μ_{it} happens to have the requisite offsetting negative correlation.[11]

The usual approach to correcting for errors-in-variables involves the use of an instrumental-variables estimator, with lagged values of the regressors or of other exogenous variables as the most frequent candidates for instruments. But the assumption of efficient markets implies that no such valid instruments can exist in our case. That is, with efficient markets, the

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[11] We might hope that OLS estimation of (13) nonetheless would provide a consistent estimate of Γ_{1i}. But this would require that $PSNEWS_t$ be uncorrelated with r_{mt} and with any other regressor that itself is correlated with r_{mt}. We found in our sample that the first of these conditions generally was satisfied, but that the second was not. In particular, we found in several cases moderate positive correlations between $PSNEWS_t$ and $AGGMNEWS_t$, and a negative correlation between $AGGMNEWS_t$ and r_{mt}.

ex post return in period t should not be systematically related to information available prior to time t. We are forced to adopt an alternative two-stage estimation procedure as follows.

We assume the validity of the so-called "market model" (see Fama, 1973); that is, we suppose that the ex post returns on all stocks are drawn from a multivariate normal distribution. This new assumption implies, but is not implied by, CAPM (Jensen, 1972). Under the normality assumption, it is always possible to write $r_{it} = \alpha_i + \beta_i r_{mt} + \epsilon_{it}$, where ϵ_{it} is orthogonal to r_{mt}. Then the constant term and the covariance term (β_i) can be consistently estimated by an OLS regression of r_{it} on r_{mt}. Once this first-stage regression has been run, we can use the estimated residuals, $\hat{\epsilon}_{it}$, as the independent variable in a regression on the various news variables. The ϵ_{it} variable can be interpreted as the "excess return" on stock i; it is the difference between the actual return to stock i in period t and that which would have been realized had the usual co-movement of the returns of that stock and the market portfolio been observed. Our second-stage equation relates the excess returns to innovations in profitability variables.

We note briefly a set of sufficient conditions for our two-step procedure to yield a consistent estimate of the parameter of interest, Γ_{1i}. First, $PSNEWS_t$ must be uncorrelated with μ_{it}. This is the small-country assumption mentioned above. While not unimpeachable, this assumption is necessary if we are to identify import competition by movements in the domestic currency price of imports (see Grossman, 1987). Next, $PSNEWS_t$ must be uncorrelated with ν_{mt}. Since the former is a sector-specific shock abroad, whereas the latter is an unobserved shock to the U.S. macro-economy, this lack of correlation seems plausible. Finally, we recognize that the unobserved ν_{mt} and the included macroeconomic news variables might be correlated. If so, then $PSNEWS_t$ must be uncorrelated with these variables, or at least with the subset of them that is correlated with ν_{mt}. We find that the sample correlations of the various PSNEWS series with all of the other news variables are relatively small, and that their correlations with the U.S. macro variables are truly negligible. Also, the other included foreign variables (AGGMNEWS, ERNEWS and WSNEWS) themselves are little correlated with the U.S. aggregative variables and can plausibly be assumed to have little correlation with ν_{mt}. We conclude that, even if ν_{mt} is not orthogonal to all of the right-hand side variables in the second-stage regression, this is not likely to be an important source of inconsistency in the estimation of Γ_{1i}.

At this point, we could apply the two-step procedure to the stock returns for each firm in our sample to obtain firm-specific measures of Γ_1. But more precise estimates are available if we impose the previously maintained assumption that all the firms in an industry share the same technology. Under this assumption, the true coefficient on each of the news variables is the same for every firm in the industry.[12] In view of the limited number of time-series observations in our sample and the considerable variability of the excess returns, we chose to impose these restrictions in our estimation of the model.[13]

By pooling the observations of returns for the different firms in an

[12] In principle, we could test the restriction that the coefficients on the news variables are equal across firms in an industry. See Gardner (1986), who conducted such a test in a similar context. However, these tests would have limited power in our small sample.

[13] In our sensitivity analysis of Section VI, we allow for the possibility that the β coefficients vary across firms, while still imposing the restriction that Γ_1 is common to firms in the industry.

industry, we created a panel data set with the special feature that all firms share common realizations of the right-hand-side variables for any time t. We treated the panel as generated by a random-effects model with time-period and idiosyncratic error components.[14] That is, we assumed that ϵ_{it}, the residual in the first-stage equation $r_{it} = \alpha + \beta r_{mt} + \epsilon_{it}$, is composed of two components. The first component, ζ_t, is time specific, but common to the firms in the industry. The second component, z_{it}, is a firm-time specific (idiosyncratic) shock, taken to be identically and independently distributed across firms and time. Similarly, we modelled η_{it}, the error term in the second-stage equation relatingthe excess return for stock i to the news variables, as the sum of a time component and an idiosyncratic component. These components represent respectively the unobserved period t shocks to productivity that similarly affect all of the firms in the industry, and that are specific to a particular firm. Note that we omit any firm-specific but time-independent components, as the efficient-market hypothesis rules out recurring (and hence predictable) shocks to the return of any given stock.

Ordinarily, to obtain efficient estimates of the random-effects model with a time-period error component, it would be necessary to use a GLS estimator to account for the heteroskedasticity of the composite error term

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[14] An altenative, but very similar procedure would be to treat the different firm equations as a system and to estimate the system by Restricted Seemingly Unrelated Regressions (RSUR). Both RSUR and the random-effects procedure allow for contemporaneous covariance between the errors for different firms. But the RSUR estimation procedure also allows the variances of the error terms in both the CAPM and excess returns equations to vary across firms, while the random-effects model imposes the restriction that these be the same. We estimated the model by RSUR and found results very similar to those reported below. (These results are available on request.) The random-effects technique is of course more efficient if the restriction it imposes is valid. A more compelling reason for selecting the random-effects alternative is that this choice facilitates calculation of the correct asymptotic standard errors of the estimates, as is discussed below.

(see Hsiao, 1986, p.34). The appropriate GLS estimator is a weighted average of the "between-period" and the "within-period" estimators, where the weights depend upon the contribution of the time-period disturbance to the overall variance of the error term (Hsiao, 1986, p.36). However, for our case where all firms in the panel are subject to a common set of observed shocks, the GLS estimator and the OLS estimator are numerically equivalent. Evidently, in the absence of any within-period variation in the right-hand-side variables, information about the correlations of the errors across firms provides no source of efficiency gain.

A final econometric point concerns the estimates of the standard errors on the coefficients. Our two-step procedure uses estimated residuals as the independent variable in the second stage. Moreover, all of the regressors at this stage are themselves estimated residuals (from the various forecasting equations). So the standard errors at the second stage should account not only for the unexplained variance in this regression, but also for the econometrician's uncertainty about the measurement of the dependent and independent variables.

Pagan (1984) has shown that when estimated residuals are used as right-hand-side variables, and the corresponding predicted values are not also included in the regression, then the estimated standard errors on the coefficients calculated by the usual least-squares formula are consistent estimates of the true standard errors. In Appendix B we show by calculations similar to his that the same is not true when, as here, the independent variable in the second stage is an estimated residual. Then the usual least-squares formula always understates the true standard errors. We have used the formula derived in the appendix to compute consistent estimates of the asymptotic standard errors of the estimated coefficients.

V. Results and Interpretation

Table 3 presents our primary results. Note that we have excluded ERNEWS from the second-stage regressions for SIC262, SIC301, SIC345, and SIC331. The inclusion of this variable in the model follows from the assumption that the exchange rate serves as a predictor of future import prices. Conversely, the model implies that if $\rho_{ej} = 0$ for all j, $\Gamma_3 = 0$. The results in Table 2 show that no lags of the exchange rate were significant in the autoregressions for p^* in four of the industries. Consequently, we dropped ERNEWS from these regressions.

The model performs admirably. In each industry, several of the news variables are found to have significant impacts on excess returns. The coefficient on PSNEWS, which theory predicts should be positive, is found indeed to be positive in all six industries. The effect of shocks to aggregate import competition on the market rate of return is less pronounced; we found $\Gamma_4 < 0$ as predicted only in four instances, and only once (in Steel Products) was the coefficient significant.

The signs of the coefficients on other variables are somewhat more difficult to interpret. The direct effect of positive PPINEWS is to increase demand for the output of each of the industries. This effect alone would lead us to expect a positive coefficient on PPINEWS, but the variable also can have indirect effects to the extent that producer prices serve as leading indicators for some of the other aggregative variables. Nonetheless, we found the coefficient on PPINEWS to be positive across the board, and significantly so in three industries. The positive coefficients on WNEWS would suggest that the industries under consideration might be more capital intensive than the average (since a positive shock to wages increases the return in these industries relative to the β-adjusted return to the market portfolio), but again the possible presence of indirect effects limits the confidence that we can place in this interpretation. A similar caveat applies to a possible interpretation that ascribes the positive coefficients on GNPNEWS to large income elasticities of demand in these industries, relative to the economy as a whole. Finally, we note that several of the industries are adversely affected by positive shocks to the price of energy; the negative coefficient in Tires and Tubes is particularly compelling. (Presumably the decline in U.S automobile production initiated by the oil-price shocks reduced the demand for domestically-produced tires).

Recall that our main objective has been to measure the sensitivity of stock market returns to import competition. We focus henceforth on the size of the estimated coefficients on PSNEWS. What do these estimates tell us about the intersectoral mobility of capital, and the likely response of shareholders in these industries to an exacerbation of import competition? To answer the first question, let us consider two extreme cases. First, suppose that $\gamma = 0$. This case corresponds to perfect, instantaneous, capital mobility, a simplifying assumption that is adopted, for example, in the Heckscher-Ohlin model. With $\gamma = 0$, adjustment is immediate, and changes in a particular import price should have no effect on the stock market returns in a small industry. Capital simply moves between industries to equate profit rates in all uses. We can test the assumption of perfect capital mobility by a one-tailed test of the null hypothesis, $H_0: \Gamma_1 = 0$, against the alternative hypothesis of imperfect mobility, where we have $H_1: \Gamma_1 > 0$. Assuming that the coefficients are approximately normally distributed, which gives a critical value for the test statistic of 1.64 for 95 percent significance, we reject the hypothesis of perfect capital mobility in five of the six industries. Only in one industry, Tires and Tubes, do the data fail to give clear evidence against the null hypothesis.

Now suppose instead that $\gamma = \infty$. This parameter value implies complete immobility of capital, as for example, in the specific-factors model. Then the size of the reduced-form parameter, Γ_1, would still depend on the forms and parameters of the profit function, $\pi(\cdot)$, and the inverse-demand function, $\phi(\cdot,\cdot)$, and also on the permanence of shocks to import prices (as revealed in the parameters ρ_{pj}). To explore the nature of this relationship, we express the reduced-form parameter Γ_1 in terms of structural parameters for the case of a Cobb-Douglas production function with capital share θ, and a constant-elasticity demand function for home goods with own-price elasticity δ and cross elasticity with respect to import prices δ^*. To determine the largest value that Γ_1 might reasonably take, we calculate this parameter for the case of a permanent shock; i.e., where $\rho_{pj} = 1$ and $\rho_{pj} = 0$ for j = 2,4.

For this extreme case of complete capital immobility, we find that $\hat{\pi}/\hat{p} = \theta$ and that $\hat{p}/\hat{p}^* = \theta\delta^*/(1-\theta+\theta\delta)$. Together these elasticities imply $\Gamma_1 = \delta^*/(1-\theta+\theta\delta)$. A typical value for θ among the industries that we study is 0.2.[15] Then, if $\delta = 4$ and $\delta^* = 2$, for example, the structural parameters would imply $\Gamma_1 = 1.25$. Alternatively, if $\delta = 2$ and $\delta^* = 1.25$, then $\Gamma_1 = 1.04$. Plausible values of the structural parameters yield maximal values for Γ_1 of

[15] We can approximate $1-\theta$ by the sum of the shares of labor compensation and materials in the value of domestic shipments. Using data for 1980 from the Annual Survey of Manufactures, we find the following values for θ: SIC242, 0.143; SIC262, 0.224; SIC301, 0.205; SIC32, 0.254; SIC345, 0.253; and SIC331, 0.097.

slightly above one, even under the counterfactual assumption that all shocks are permanent. Evidently, estimated values for Γ_1 at or near 1.0 would imply a high degree of capital specificity.[16]

Table 3 reports estimated coefficients on PSNEWS of greater than 0.8 for four of the six industries studied. In view of the fact that the actual shocks to import prices in these industries, while relatively long-lasting, are far from permanent, a striking conclusion emerges. At least in four of the six industries, namely Paper, Nuts and Bolts, Ceramics, and Steel, the prospects for intersectoral movements of capital in response to changes in profit opportunities seem to be quite limited indeed. Trade models that assume complete immobility of capital may come much closer to capturing the reality for U.S. industries than do ones that assume instead perfect mobility.

Finally, we remark on the quantitative significance of import competition as a cause of fluctuations in stock market returns. Of course, stock returns are highly volatile, and unanticipated movements in import prices explain but a small fraction of this variability. Nonetheless, the distributional implications of terms of trade changes for owners of capital invested in import-competing industries are hardly negligible. Table 2 reports the

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[16] We note two caveats to this remark. First, we have implicitly assumed for purposes of these calculations that firms are 100 percent equity financed. Since variations in profit affect equity values more than they do debt values, the stock returns of a firm that is leveraged (partially financed by debt) will show greater sensitivity to shocks than one that is not. Second, we have assumed that the firms have all of their capital invested in one industry. In fact, most firms produce goods in more than one SIC category. Diversification reduces the sensitivity of stock returns to shocks in any one sector. Allowing for these considerations, and assuming that profit shocks affect <u>only</u> the value of equity, we find that with complete capital immobility and permanent price shocks, $\Gamma_1 = y\delta^*/(1\text{-}z)(\theta\delta+1\text{-}\theta)$. Here y is the fraction of the firm's value that it derives from profits in the import-competing industry and z is the share of debt in the total value of the firm. Notice that the two omitted factors have offsetting implications for the maximal value of Γ_1.

standard deviation of the unanticipated shocks to import prices in each of the industries. A one standard deviation shortfall of the import price in any quarter relative to its predicted level can mean capital losses for owners of shares of firms in the Paper, Lumber, Ceramics, Steel, and Nuts and Bolts industries of between 1.4 and 3.0 percent.

VI. Sensitivity Analysis

Leamer (1983) has argued persuasively that because any econometric analysis involves numerous debatable decisions, findings cannot be convincing unless they are shown to be robust. In this section, we explore the sensitivity of our coefficient estimates to changes in several of the modelling decisions that were made along the way.

One area where theory offers little guidance concerns the way in which investors form their expectations. A rational investor will use all available information, provided that the benefits of doing so do not exceed the collection and computation costs. Still, the econometrician cannot observe the investor's information set, nor can he know what information actually is used and in what manner. Thus, any specification of expectations should be viewed as doubtful (to use Leamer's terminology) in empirical analysis.

In our base-case estimation, we formed PSNEWS, the unanticipated component of the current period import price, by a procedure of nested hypothesis testing. We assumed, in effect, that investors used exactly as many lags of foreign wages, exchange rates, and the import price itself in predicting detrended import prices as were shown to be statistically significant in a multivariate autoregression. Other specifications clearly have equal claim to plausibility. Here we consider three alternatives. We formed PSNEWSA as the residual of a regression of $\tilde{p}^*$ on itself lagged once, and on four lags each of $\tilde{w}^*$ and $\tilde{e}$. Under this specification, investors are assumed to use foreign wages and exchange rates as predictors regardless of the statistical significance of these variables in the autoregression. PSNEWSB is the residual from a simple, first-order autoregression for $\tilde{p}^*$. We motivate this specification with reference to the fact that, across all of the industries in our study, the first lag of $\tilde{p}^*$ explains far more of the variance of $\tilde{p}^*$ than any of the other variables in the autoregressions. Moreover, a first-order autoregression certainly is the simplest procedure for investors to implement. Finally, we formed PSNEWSC by taking first differences of the log of the (non-detrended) import price. Until now, our various methods for generating PSNEWS all have relied upon deterministic techniques for removing the trend from the import-price series. If the (logs of) import prices actually were to follow a random walk, then deterministic detrending would introduce spurious cyclicality into the series (see Nelson and Kang, 1981), and biased estimates of the autoregressions would result. The correct measure of innovations in import prices in this case would be PSNEWSC.[17]

Table 4 reports the estimated coefficient on import price news (Γ_1) for each of the alternative specifications of how expectations about import prices are formed. Broadly speaking, the estimates of Γ_1 seem to be robust with respect to alternative specifications of PSNEWS. Most of the estimates in the second through fourth columns of Table 4 are within one standard error of the

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[17] A second advantage of the log-difference specification is that it does not use information from "future" years in generating expectations about next quarter's variables. Strictly speaking, the other procedures require an implicit assumption that the (time-invariant) processes for the exogenous variables are known to investors from the outset, and the econometrician estimates the autoregressions to learn what the investors already know.

corresponding estimates for the base case. The conclusion that capital is nearly completely immobile does not, perhaps, emerge quite as forcefully in Table 4 as it does in Table 3. When the alternative measures of PSNEWS are used, only three of the six industries are consistently found to have estimated coefficients on this variable in excess of 0.6. But the evidence against perfect intersectoral mobility of capital remains strong and convincing. In fact, when PSNEWSB is used as the measure of unanticipated shocks to import prices, the hypothesis of perfect capital mobility is rejected for all six industries.

Not only is the investors' information set unobserveable to the econometrician, but so too is the timing of the arrival of information. To this point, we have assumed that investors learn the realizations of all variables in the current period. Another possibility is that some or all of these variables enter the information set with a lag of one quarter. Then current-period excess returns would respond to news lagged once. We experimented with several specifications in which lagged news was entered either separately, or in combination with contemporaneous news. In no case was the coefficient on a lagged news variable statistically significant, and the inclusion of the lagged variables had little effect on the estimated coefficients for current-period news.

Finally, one might question the restrictions imposed on the data by the CAPM specification. In particular, our estimation has presumed that the relationship between individual stock returns and the market return remained constant throughout the sample period, and that all firms in an industry experienced equal exposure to risk. We relaxed these assumptions for the estimation reported in Table 5. The first column repeats our base-case estimates of Γ_1. Next, we removed the restriction that β is constant over time, allowing instead for β to follow a quadratic time trend. That is, we assumed that we could approximate $\beta(t)$ by $\beta_1 + \beta_2 t + \beta_3 t^2$.[18] For column (2), we continued to assume that firms in an industry share common values of $\beta(t)$. The estimates in column (3) were generated by again imposing time-invariance for β, but now relaxing the restriction that this value be common to all firms.[19] Finally, column (4) shows the estimated coefficients on PSNEWS when both the across-time and across-firm restrictions are removed simultaneously.

Once again, the estimates of Γ_1 prove reasonably robust to variations in specification. We continue to find substantial evidence contradicting the assumption of perfect capital mobility in five of the six industries studied, and considerable support for the hypothesis of complete capital specificity in four of these cases.

VII. Concluding Remarks

In this paper, we have developed a method for assessing the sensitivity of stock market returns to variations in import competition and other exogenous variables affecting firms' profit streams. The method relies on identification of the unanticipated component of variables that enter the reduced-form profit function. These innovations determine the excess of

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[18] Note that equation (13) implies that if β varies over time, so to do the constant in the first-stage regression and the coefficients Γ_j, for $j = 4,9$, in the second-stage regression.

[19] When β is allowed to vary across firms, then so too should the coefficients Γ_j for $j = 4,9$ in the second-stage regression be alloed to do so. Thus, the random-effects model, which imposes that these coefficients be the same, no longer is appropriate. Instead, we generated the estimates reported in columns (3) and (4) by Restricted Seemingly Unrelated Regressions, where only the coefficients on industry-specific variables were constrained to be the same across firms in an industry.

realized returns on a firm's equity over the *ex ante* expected returns. The estimated reduced-form coefficients can be given a structural interpretation by solving the reduced form for certain constellations of the structural parameters.

By applying the method to data for six U.S. industries, we were able to test the hypothesis of perfect intersectoral capital mobility against the alternative hypothesis of imperfect mobility. The data reject the assumption of perfect mobility at the 95 percent level of significance in five of the six cases. Furthermore, in four industries, Paper, Nuts and Bolts, Ceramics, and Steel, the size of the estimated coefficient on import price news is quite consistent with the opposite extreme hypothesis of complete specificity of the industry capital stock. The estimates indicate that a one standard deviation shock to the expected import price creates substantial capital gains or losses for shareholders in at least five of the six industries.

Of course, the method developed here warrants further refinement. More sophisticated procedures for identifying the unanticipated shocks to the profitability variables might improve upon the precision of the estimates. Also, recent advances in finance theory might allow a better decomposition of realized returns into expected and extra-normal components. We hope that the encouraging results reported here will spur additional empirical research on the distributional effects of international competition and trade policy.

APPENDIX

A. Capital Mobility and the Sensitivity of Stock Returns to Import Prices

We derive formally the relationship between capital mobility, as measured by the parameter γ in the model of Section II, and the sensitivity of stock market returns to changes in import prices.

Following a permanent, unanticipated change in the import price, the industry adjusts to a new, steady-state equilibrium. Let ψ_i denote the slope of the saddlepath for firm i. From equations (2) and (3), we have

$$\psi_i = \frac{d\lambda}{dK_i} = \frac{\gamma(r_f\lambda - \pi) - \gamma^2 I_i^2/2K_i^2}{(\lambda - 1)K_i}$$

$$= \frac{\gamma(r_f\lambda - \pi)}{(\lambda - 1)K_i} - \frac{(\lambda - 1)}{2K_i} \quad . \tag{A1}$$

Now consider how the slope of the saddlepath varies with a change in γ:

$$\frac{d\psi_i}{d\gamma} = \frac{r_f\lambda - \pi}{(\lambda - 1)K_i}$$

$$= \frac{\psi_i}{\gamma} + \frac{(\lambda - 1)}{2K_i} \tag{A2}$$

Since $\psi_i \leq 0$, $\lambda < 1 \Rightarrow d\psi_i/d\gamma < 0$. That is, for points to the right of the steady-state equilibrium, where dis-investment takes place over time, an increase in the adjustment-cost parameter unambiguously tilts the saddlepath in a clockwise direction. This implies a greater sensitivity of share values to declining import prices, as we shall see shortly.

What about the case where $\lambda > 1$, so that the industry capital stock must rise over time? It is impossible to say, in general, how an increase in γ tilts the saddlepath at some arbitrary point to the left of the steady-state equilibrium. However, we can establish that $d\psi_i/d\gamma < 0$ at least in the neighborhood of a steady-state equilibrium. Taking the limit of (A2) as $K_i \to \bar{K}_i$, and denoting the limiting slope by $\bar{\psi}_i$, we find

$$\lim_{K_i \to \bar{K}_i} \frac{d\psi_i}{d\gamma} = \frac{\bar{\psi}_i}{\gamma} ,$$

where, from (A1),

$$\bar{\psi}_i = \lim_{K_i \to \bar{K}_i} \frac{\gamma(r_f \lambda - \pi)}{(\lambda - 1)K_i} . \tag{A4}$$

In a steady state, the right-hand side of (A4) is 0/0. Applying L'Hospital's rule, we have

$$\bar{\psi}_i = \frac{\gamma r_f \bar{\psi}_i - \gamma \pi_K}{K_i \bar{\psi}_i} . \tag{A5}$$

where π_K represents the change in profits as we increase the firm's capital, thereby moving closer to the steady state. Since all firms in the industry follow similar investment profiles, we know that the industry capital stock grows in proportion to the growth in K_i, hence $\pi_K < 0$. Finally, we solve the quadratic equation in (A5), and select the appropriate (negative) root for stability, which gives

$$\bar{\psi}_i = \frac{\gamma r_f - [\gamma^2 r_f^2 - 4\gamma K_i \pi_K]^{1/2}}{2K_i} . \tag{A6}$$

From (A6) we see that $\gamma > 0 \Rightarrow \bar{\psi}_i < 0$, hence $d\psi_i/d\gamma < 0$ near the steady state.

Having established that an increase in γ makes the saddlepath steeper near the steady state, we can apply Hayashi's theorem to derive the implications for stock prices. Consider an industry in steady-state equilibrium that experiences a sudden jump in the import price. A steeper saddlepath to the new steady state means that, immediately following the increase (decrease) in the import price, with K still at its initial level, the shadow value of the installed capital is larger (smaller). But the fact that marginal q equals average q implies $\lambda = V_i/K_i$. We conclude that a firm of given size will experience a larger increase (decrease) in its stock price following an increase (decrease) in the import price, when capital in the industry is less mobile.

B. Correction Factors for the Standard Errors

In this appendix, we derive the correction factors for the estimated standard errors. These corrections are needed to account for the fact that the two-stage procedure uses estimated residuals from the first stage as the independent variable in the second-stage regression.

Consider the following two equation model:

$$y = x\beta + \epsilon \tag{B1}$$

$$\epsilon = z\alpha + \eta \tag{B2}$$

Suppose that OLS estimation of (B1) yields estimated residuals $\hat{\epsilon}$, and that these estimated residuals are used in place of ϵ in OLS estimation of equation (B2). Then the true (asymptotic) variance of the estimated vector of parameters $\hat{\alpha}$ is given by

$$\lim_{T\to\infty} E\, T(\hat{\alpha} - \alpha)(\hat{\alpha} - \alpha)' = \text{plim}\ T[\sigma_\eta^2(z'z)^{-1} + 2\sigma_{\epsilon\eta}(z'z)^{-1}z'x(x'x)^{-1}x'z(z'z)^{-1} + \sigma_\epsilon^2(z'z)^{-1}z'x(x'x)^{-1}x'z(z'z)^{-1}] \qquad \text{(B3)}$$

and since $\sigma_{\epsilon\eta} = E\ \epsilon\eta' = E\ \eta\eta' + \alpha E\ z'\eta = \sigma_\eta^2$, we can rewrite (B3) as

$$\text{asy cov}\ \hat{\alpha} = \text{plim}\ T[\sigma_\eta^2(z'z)^{-1} + (2\sigma_\eta^2 + \sigma_\epsilon^2)(z'z)^{-1}z'x(x'x)^{-1}x'z(z'z)^{-1}]$$

How does this true asymptotic covariance matrix compare to the covariance of the estimates as reported by OLS? Let $\hat{\eta}$ be the estimated residuals from the second regression, and define $\xi \equiv \hat{\epsilon} - \epsilon$. Then $\hat{\epsilon} = z\alpha + \eta + \xi$ and

$$\hat{\eta} = [I - z(z'z)^{-1}z'](\eta + \xi)$$

Forming the sum of squared residuals, we have

$\hat{\eta}'\hat{\eta} = (\eta + \xi)'[I - z(z'z)^{-1}z'](\eta + \xi)$. Now since:

$$\text{plim}\ (1/T)\eta'\xi = \text{plim}\ (1/T)\eta'x \cdot \text{plim}\ T(x'x)^{-1} \cdot \text{plim}\ (1/T)x'\epsilon = 0\ ;$$

$$\text{plim}\ (1/T)z'\eta = 0\ ;$$

$$\text{plim}\ (1/T)z'\xi = \text{plim}\ (1/T)z'x \cdot \text{plim}\ T(x'x)^{-1} \cdot \text{plim}\ (1/T)x'\epsilon = 0\ ;$$

and

$$\text{plim } (1/T)\xi'\xi = \text{plim } (1/T)e'x \cdot \text{plim } T(x'x)^{-1} \cdot \text{plim } (1/T)x'\epsilon = 0 \ ,$$

we find $\text{plim } (1/T)\hat{\eta}'\hat{\eta} = \sigma_\eta^2$. Thus, the familiar OLS formula gives the covariance matrix for $\hat{\alpha}$ as something that converges in probability to $\text{plim } T(z'z)^{-1}\sigma_\eta^2$. This understates the true asymptotic covariances by $F \equiv (2\sigma_\eta^2 + \sigma_\epsilon^2)\cdot\text{plim } [T(z'z)^{-1}z'x(x'x)^{-1}x'z(z'z)^{-1}]$.

We obtained a consistent estimate of the aymptotic covariance matrix of the parameter estimates by adding the matrix F to the estimated covariance matrix as calculated by the usual least-squares formula. In Table B1 we have tabulated the ratios for our base case of the corrected standard errors to those computed without accounting for the two-stage procedure. As can be seen, the corrections factors generally are small, and those for the standard errors of the estimates of Γ_1 never exceed eight percent. This is not surprising in the light of the small correlations between r_{mt} and the various news variables that we find in our sample.

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Figure 1

Adjustment of Capital Stock

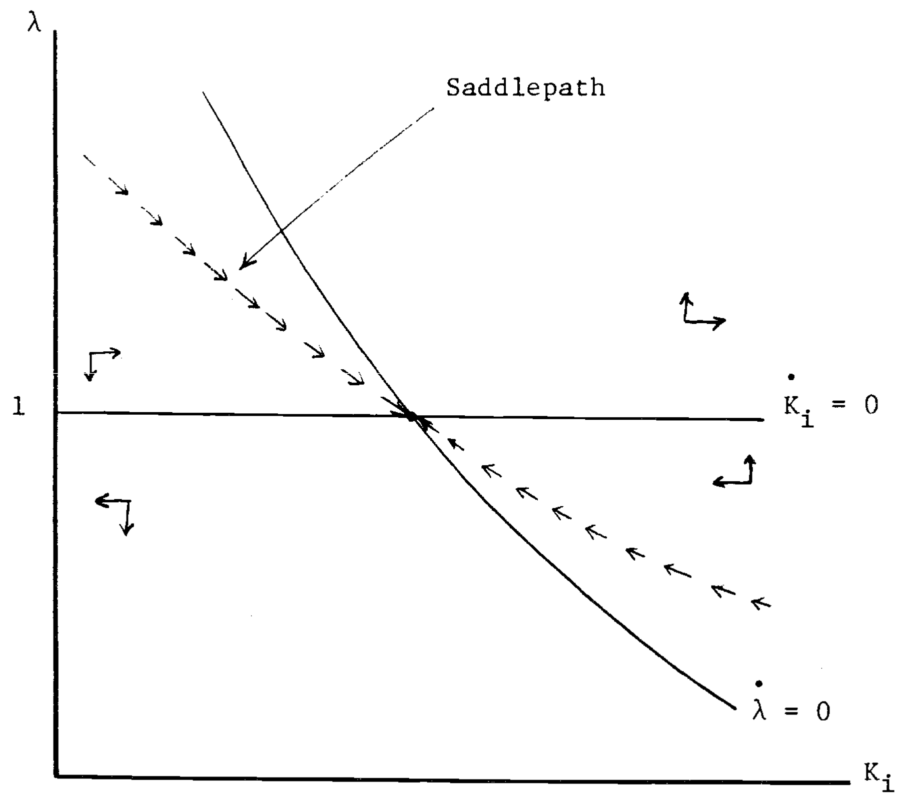

Figure 2

Unanticipated Increase in Import Price at Time T

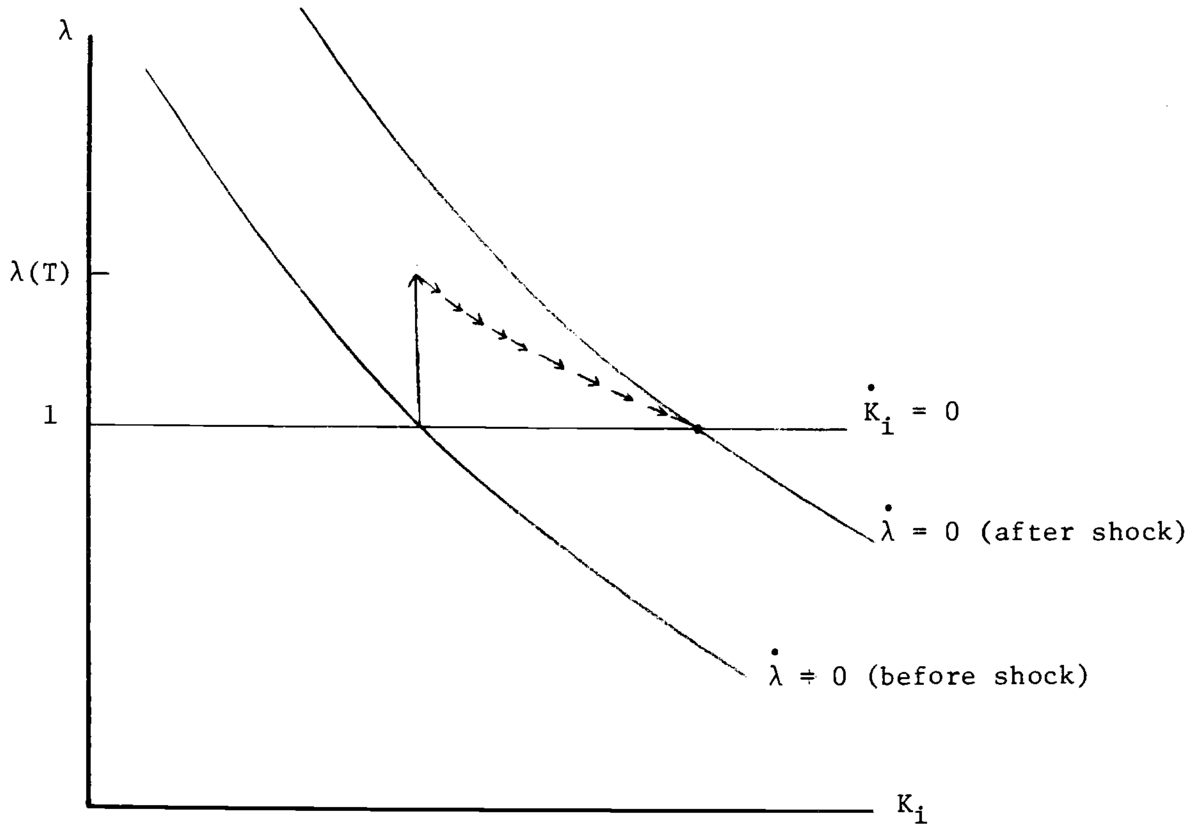

Table 1

Variable Definitions and SIC Categories

Variables

r_{it} -- The stock market return (dividends plus capital gains) to firm i in period t.

r_{mt} -- The return (dividends plus capital gains) on a value-weighted portfolio of all NYSE stocks.

PSNEWS -- The news to an industry specific (SIC-based) import price index.

AGGMNEWS -- The news to an aggregate import price index for the United States

PENEWS -- The news to an index of energy prices for the United States.

WNEWS -- The news to an index of non-agricultural wages in the United States.

GNPNEWS -- The news to U.S. Gross National Product.

PPINEWS -- The news to a producer price index for the United States.

WSNEWS -- The news to an index of foreign wages for suppliers of a particular import good to the United States.

ERNEWS -- The news to an index of bilateral exchange rates for suppliers of a particular import good to the United States.

SIC Code Groups

SIC 262 -- Paper Mill Products.

SIC 242 -- Sawmill and Planing Mill Products.

SIC 301 -- Tires and Inner Tubes.

SIC 345 -- Nuts, Screws, Rivets, etc. of Base Metal.

SIC 32 -- Stone, Clay, Glass, and Concrete Products.

SIC 331 -- Steel and Rolling Mill Products.

Note: See the text for details of the construction of the various "news" variables. Original sources of the data are provided in Section IV.B and in a separate Data Appendix, available upon request.

Table 2

Specification Tests for Import Price News Variable

(See interpretative note below.)

Lags	Restriction	SIC262	SIC242	SIC301	SIC345	SIC32	SIC331
$\tilde{p}^*$: 4 $\tilde{w}^*$: 8 $\tilde{e}$: 8	$\tilde{p}^*$:2-4= 0	.939	.391	.251	.004	.694	.087
	$\tilde{w}^*$:5-8= 0	.548	.317	.283	.458	.644	.358
	$\tilde{e}$:5-8= 0	.711	.090	.734	.875	.937	.751
$\tilde{p}^*$: 1,4 $\tilde{w}^*$: 4,8 $\tilde{e}$: 4,8	$\tilde{w}^*$:2-4= 0	.114	.054	.026	.661	.041	.078
	$\tilde{e}$:2-4= 0	.384		.895	.254	.191	.420
$\tilde{p}^*$: 1,4 $\tilde{w}^*$: 1,4,8 $\tilde{e}$: 1,4,8	$\tilde{w}^*$: 1= 0	.031			.004		
	$\tilde{e}$: 1= 0	.153		.314	.342	.001	.353
Final lag structure on news ($\tilde{p}^*$,$\tilde{w}^*$,$\tilde{e}$)		(1,1,0)	(1,4,8)	(1,4,0)	(4,1,0)	(1,4,1)	(4,4,0)
Estimation Period		1974:2 to 1986:4	1974:3 to 1986:4	1976:3 to 1986:3	1975:2 to 1986:3	1974:2 to 1986:4	1975:1 to 1986:3
Stand.Dev. of PSNEWS		.022	.048	.012	.018	.017	.034

Variable Definitions:

- $\tilde{p}^*$ --The deterministically detrended, deseasonalized log of the import price index for a specific (SIC-based) industry
- $\tilde{w}^*$ --Weighted average of deterministically detrended, deseasonalized logs of foreign wages in foreign currency, where weights are import shares in the SIC group.
- $\tilde{e}$ --Weighted average of deterministically detrended, deseasonalized logs of exchange rates, where weights are import shares in the SIC group.

Note on interpretation: Cell entries are significance levels for F-tests. Each F-test is a test of zero-restrictions on the lag structure of the autoregression for import prices. A cell value <u>less than</u> .10 indicates rejection at the 90 percent confidence level. Failure to reject causes elimination of the relevant lags before proceeding to next level.

Table 3

Estimation of Random-Effects Model with Time Components
(Base Case Results)

Step 1: $r_{it} = \alpha + \beta r_{mt} + \epsilon_{it}$

Step 2: $\hat{\epsilon}_{it} = \Gamma_1 PSNEWS_t + \Gamma_2 WSNEWS_t + \Gamma_3 ERNEWS_t + \Gamma_4 AGGMNEWS_t + \Gamma_5 PENEWS_t + \Gamma_6 WNEWS_t + \Gamma_7 GNPNEWS_t + \Gamma_8 PPINEWS_t + \Gamma_9 MSNEWS_t + \eta_{it}$

	SIC262	SIC242	SIC301	SIC345	SIC32	SIC331
Step 1 Results						
r_{mt}	1.116 *	1.502 *	1.439 *	1.358 *	1.080 *	1.081 *
	(.063)	(.083)	(.128)	(.206)	(.058)	(.071)
Step 2 Results						
PSNEWS	1.222 *	.434 *	.185	1.548 †	.804 *	.893 *
	(.295)	(.171)	(.926)	(.920)	(.371)	(.170)
WSNEWS	-.719	-.147	-1.549	-2.112	1.474 *	2.115 *
	(.554)	(.718)	(2.02)	(1.68)	(.714)	(.737)
ERNEWS	---	.333	---	---	.043	---
		(.479)			(.342)	
AGGMNEWS	-.0062	.079	.012	-.072	-.131	-.447 *
	(.180)	(.246)	(.384)	(.563)	(.161)	(.190)
PENEWS	-.109	-.800 *	-.596	-.291	-.720 *	-.132
	(.233)	(.298)	(.365)	(.632)	(.197)	(.268)
WNEWS	4.158 *	3.414 *	4.527 *	.416	2.682 *	.004
	(.831)	(1.10)	(1.36)	(2.32)	(.883)	(.966)
GNPNEWS	2.752 *	.024	1.923	4.234 †	.717	1.549 *
	(.778)	(1.09)	(1.36)	(2.16)	(.696)	(.785)
PPINEWS	.754	3.052 *	2.713 *	1.932	1.953 *	1.377
	(.757)	(1.03)	(1.36)	(2.35)	(.704)	(.883)
MSNEWS	2.272 *	1.361	1.464	3.165	2.445 *	1.116
	(.883)	(1.10)	(1.39)	(2.41)	(.830)	(.968)
# of firms in SIC group	9	5	7	2	16	16
Estimation Period	1974:2 to 1985:4	1974:3 to 1986:3	1976:3 to 1986:3	1975:2 to 1986:3	1974:2 to 1986:3	1975:1 to 1984:2

Note: Standard errors are in parentheses. An asterisk (dagger) indicates that coefficient is significantly different from 0 at the 95% (90%) confidence level.

Table 4

Alternative Specifications of PSNEWS

Coefficient on industry-specific import price news (Γ_1)

	PSNEWS	PSNEWSA	PSNEWSB	PSNEWSC
SIC 262	1.222 *	1.305 *	1.201 *	1.200 *
	(.295)	(.326)	(.281)	(.249)
SIC 242	.434 *	.362 *	.305 *	.312 *
	(.171)	(.163)	(.138)	(.129)
SIC 301	.185	-.091	1.667 *	-1.025
	(.926)	(.974)	(.902)	(.872)
SIC345	1.548 *	.827	1.219 *	1.116 *
	(.920)	(.788)	(.592)	(.616)
SIC32	.804 *	1.136 *	.832 *	.996 *
	(.371)	(.388)	(.290)	(.299)
SIC331	.893 *	.606 *	.583 *	.332 *
	(.170)	(.160)	(.145)	(.124)

Variable Definitions:

PSNEWS --The specification resulting from testing the significance of lagged values of foreign wages, import prices, and exchange rates as summarized in Table 2. This variable is used in the base case estimation.

PSNEWSA --The import price news variable for an SIC category resulting from an autoregression of the deterministically detrended, deseasonalized log of the SIC import price index on one lag of $\tilde{p}*$, and four lags each of $\tilde{w}*$ and $\tilde{e}$.

PSNEWSB --The import price news variable for an SIC category resulting from a first-order autoregression of deterministically detrended, deseasonalized, log of the SIC import price index.

PSNEWSC --The import price news variable for an SIC category resulting from first-differencing the log of the SIC import price index.

Note: Specification of the model is the same as for base case. Estimation using PSNEWSA and PSNEWSB always excludes WSNEWS and ERNEWS from step 2. Estimation using PSNEWSC always includes these variables. An asterisk indicates rejection in a one-tailed test of the hypothesis of perfect capital mobility at the 95% confidence level. Standard errors are in parentheses.

Table 5

Alternative Specifications of CAPM

CAPM: $r_{it} = \alpha_{0i} + \alpha_{1i}t + \alpha_{2i}t^2 + \beta_{1i}r_{mt} + \beta_{2i}(r_{mt}*t) + \beta_{3i}(r_{mt}*t^2) + \epsilon_{it}$

Restriction Set #1: $\alpha_{0i} = \alpha_0$; $\alpha_{1i} = \alpha_{2i} = 0$; $\beta_{1i} = \beta_1$; $\beta_{2i} = \beta_{3i} = 0$.
Restriction Set #2: $\alpha_{0i} = \alpha_0$; $\alpha_{1i} = \alpha_1$; $\alpha_{2i} = \alpha_2$; $\beta_{1i} = \beta_1$; $\beta_{2i} = \beta_2$; $\beta_{3i} = \beta_3$.
Restriction Set #3: $\alpha_{1i} = \alpha_{2i} = 0$; $\beta_{2i} = \beta_{3i} = 0$.
Restriction Set #4: No restrictions.

Coefficient on PSNEWS
for Alternative Specifications of CAPM

	Restriction Set on CAPM			
SIC Category	(1)	(2)	(3)	(4)
SIC262	1.222 *	1.090 *	1.414 *	1.832 *
	(.295)	(.381)	(.338)	(.364)
SIC242	.434 *	.547 *	.340 *	.319 *
	(.171)	(.198)	(.203)	(.198)
SIC301	.185	.390	-.297	-1.008
	(.926)	(1.008)	(.826)	(.713)
SIC345	1.548 *	2.317 *	.976	2.114 *
	(.658)	(1.393)	(.613)	(.595)
SIC32	.804 *	.915 *	.849 *	.672 *
	(.371)	(.448)	(.335)	(.447)
SIC331	.893 *	.661 *	1.004 *	.881 *
	(.170)	(.199)	(.205)	(.172)

Note: Columns (1) and (2) estimated as random-effects-time-components model. Columns (3) and (4) estimated by Restricted Seemingly Unrelated Regressions. An asterisk indicates rejection in a one-tailed test of the hypothesis of perfect capital mobility at the 95% confidence level. Standard errors are in parentheses. For reasons of cumputational complexity, those in columns (3) and (4) have not been adjusted to account for the two-step procedure.