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Abstract

This paper estimates the long run elasticity of the demand for fixed nonresidential capital (both equipment and structures) to changes in its user cost using a quarterly panel of two-digit manufacturing data from South Africa from 1970 to 2001. Using a difference specification that does not rely on cointegration, we find highly significant estimates of the user cost elasticity on the order of -0.80. These estimates contrast sharply with many previous studies that obtained small and/or statistically insignificant estimates of the user cost elasticity using U.S. data. This discrepancy may owe to the possibility that the capital demand curve is better identified in a small open economy because shocks to capital supply are more likely to be exogenous. The economic embargo imposed on South Africa from 1985 to 1993 forced its economy to become more closed and therefore provides a unique natural experiment to assess this conjecture. Estimates of the user cost elasticity over this period are small and statistically insignificant, similar to the findings of previous studies where the user cost was likely endogenous. These findings underscore the importance of identification in estimating the user cost elasticity of capital demand.

Keywords: user cost elasticity; fixed investment; capital accumulation; price of capital; interest rate; tax policy

JEL classifications: E22, E44, E62, C23

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1 Introduction

Understanding the determinants of capital accumulation is essential to understanding business cycles and economic growth, and therefore to effective formulation of economic policy. It is no surprise that estimating the response of capital demand to changes in the user cost of capital (the user cost elasticity) has been one of the most widely researched areas in empirical macroeconomics. Despite the voluminous research on the subject, the results remain somewhat inconclusive. Estimates using aggregate time series data have found an elasticity that is statistically insignificant or, counter-intuitively, positive in sharp contrast to theory based on the neoclassical framework.[1]

According to the neoclassical theory, both the capital stock and the user cost of capital are determined by demand and supply equilibrium that equates the marginal product of capital services to its marginal opportunity cost. In order to obtain reliable estimates of the user cost elasticity of the capital demand curve, econometricians must be able to isolate exogenous movements in the supply curve for capital. This is particularly challenging in a large open economy like the United States, where the demand and supply of capital services are jointly determined. Estimates that fail to account for this simultaneity are likely to be biased.[2] The presence of capital adjustment costs also complicates the estimation process. These costs not only prolong the response of capital to a given change in the user cost, but also cause the magnitude of the response to be closely related to both the size and the anticipated persistence of these shocks (Tevlin and Whelan [2003]). This places an emphasis on obtaining long-run estimates that cut through the noise from transitory changes in investment fundamentals.

This study expands on the insight by Schaller [2006] that in a small open economy, movements in the supply of capital services are more likely to be exogenous because the country is a price taker in the world markets for financial capital and investment goods. We use a quarterly panel of two-digit manufacturing data from South Africa and use a regression technique that accounts for internal adjustment costs by focusing on the long run response of capital. Since South Africa

[1] The alternative theory is the well-known "accelerator model" that suggests a demand for capital is driven by changes in output. According to this view, price variables including the user cost play little or no role in the demand for capital once the effect due to changes in output are taken into account. See Chirinko [1993] and Hassett and Hubbard [2002] for surveys.

[2] In principle, the simultaneity problem can be addressed by using instrumental variables. As noted by Hassett and Hubbard [2002], conventional instrumental variables, such as lagged endogenous variables or sales-to-capital ratios, have not proven successful.

is a small open economy, its domestic demand for capital has a limited effect on interest rates and the price of capital goods, and hence on the user cost, allowing a better estimate of the user cost elasticity.[3] Our quarterly panel helps both to emphasize the low frequency movements in the data that are important for capital accumulation and to limit the potential for small-sample bias, which is heightened in this setting due to the serial persistence properties of the data.

We find an estimated user cost elasticity of capital demand in the vicinity of -0.80 that is highly statistically significant and reasonably close the value embedded in the Cobb-Douglas production function that is often assumed by researchers. To our knowledge, the only other studies that have found such a large and significant elasticity are Caballero [1994] and Schaller [2006], who assume a cointegrating relationship between capital, output and the user cost and whose headline estimates of the user cost elasticity use measures of business fixed capital that exclude structures.[4] Our study is the first to document such a large elasticity using data for a measure of capital that includes both equipment and structures, and that employs a lagged-difference approach rather than a cointegration technique. We also obtain similar estimates when we use a panel cointegration approach, although our panel cointegration tests suggest that the cointegrating relationship in Caballero [1994] and Schaller [2006] may not be very robust for the South Africa data.

An additional feature of the South African economy exploited in our study is the unique reversion toward autarky that the country experienced beginning around 1985 until early 1994. During this period, the world imposed economic sanctions on South Africa to put pressure on its apartheid regime—a political system that granted different rights to citizens based on race. As a result of the embargo, several foreign public and private entities operating in South Africa decided to disinvest and/or stop making new investments or reinvestments of earnings in the country.[5] In addition

[3]Even though prices and interest rates were determined in world markets, it is possible that the user cost may have moved endogenously owing to changes in corporate tax rates. However, corporate tax rates were relatively constant over much of our sample period. Although tax rates were lowered in the 1990s, these changes were largely motivated by the need to keep rates competitive internationally. See the Katz Commission Third Interim Report of the Commission of Inquiry in to Certain Aspects of the Tax Structure of South Africa, The Government Printer, Pretoria (1995).

[4]Structures may not be a negligable omission from these estimates. For example, according to estimates by the Bureau of Economic Analysis, over the post-war period from 1946 to 2005, structures accounted for an average of almost two-thirds of the total stock of private nonresidential fixed capital in the United States (current value terms), and an average of about one-third of the nominal value of business fixed investment.

[5]The International Monetary Fund estimated that the embargo cost South Africa $8 billion in foregone foreign investment between 1985 and 1991, about 3 percent of the country's cumulative GDP from 1985 to 1991. The U.S. General Accounting Office (GAO) estimated that $10.8 billion flowed out of South Africa from January 1985 through

to these restrictions on capital flows, several countries also restricted or banned trade with South Africa. These restrictions had a meaningful effect on the country's international trade flows. South Africa's trade-to-GDP ratio dropped from an average of 23 percent during the pre-embargo period to an average of 19 percent during the embargo, then rose back to an average of 25 percent after the embargo was lifted. The country's current account balance, shown in Figure 1, also follows a pattern consistent with these restrictions.[6] Before 1985, the country registered current account deficits that averaged 2 percent of GDP. However, when economic sanctions intensified between 1985 and 1993, the current account balance swung to a surplus that averaged about 2.4 percent of GDP. After 1993, the current account reversed again to a deficit as the economic sanctions were lifted and the country re-integrated into the world economy.[7]

South Africa's limited access to world markets when the embargo was in effect suggests that the variables that determine the user cost, such as interest rates and the relative price of capital goods, became much more influenced by domestic factors. We exploit this 'natural' experiment to test whether the simultaneity problem has a meaningful effect on our capital elasticity estimates. We find that controlling for the data from the embargo period leads to a statistically significant increase in the absolute value of the user cost elasticity, and that the estimated elasticity during the embargo period is considerably smaller, and, in some cases, close to zero. This finding may help explain why many previous studies that employ data from large economies have had difficulty finding estimates of the capital user cost elasticity that are statistically significant and of the correct sign.

June 1989, including $3.7 billion in loan repayments to banks, $7.1 billion in other debt repayments and capital flight (GAO 1990, 12, 17). Similarly, Trust Bank (a South African commercial bank) calculated that the country had forgone nearly $14 billion in loans and direct investments between 1985 and 1990 in comparison to what loans and direct investments would have been had money flowed in at the rates that had prevailed before 1985 (The Economist, 10 February 1990, 69).

[6] Detailed historical accounts of the economic embargo and the disinvestment are available on the Institute for International Economics website at the following addresses: http://www.iie.com/research/topics/sanctions/southafrica3.htm#economic http://www.iie.com/research/topics/sanctions/southafrica.htm#chronology

[7] See Coulibaly [2005]

2 Theoretical Framework

We assume that a forward-looking representative firm in each industry chooses a quantity of capital that maximizes its market value in the face of capital adjustment costs. This choice balances the costs of adjustment against the costs associated with departing from the capital that it would hold in a frictionless world. The *frictionless capital stock* is determined by the neoclassical investment fundamentals, and takes the form:

$$k_{i,t}^{*} = \beta y_{i,t} - \sigma u_{i,t} + (\sigma - 1)\, a_{i,t} \qquad \text{for industries} \qquad i = 1, ..., N, \tag{1}$$

where $y_{i,t}$ and $u_{i,t}$ are the log of output and the log of the user cost for frictionless capital in industry i.[8] The variable $a_{i,t}$ is the log of the level of technology, an important fundamental for capital holdings that we assume is known by firms in industry i but is not observed by econometricians. The user cost for the frictionless capital stock in each industry is given by:

$$U_{i,t} = \frac{\overline{p}_{i,t-1}^{k}}{p_{it}(1-\tau_t)} \left[r_{t-1} + \zeta + \delta_{i,t} - \frac{E_{t-1}\left[\Delta \overline{p}_{i,t}^{k}\right]}{\overline{p}_{i,t-1}^{k}} \right] \tag{2}$$

where τ_t is the corporate tax rate, $\overline{p}_{i,t}^{k}$ is the price of capital goods net of the present value of all future tax shields from depreciation allowances, r_t is the interest rate, ζ is a risk premium for capital, $\delta_{i,t}$ is the depreciation rate in industry i, and $p_{i,t}$ is the price of output in industry i.[9]

We follow Hall and Jorgenson [1967] and many others by assuming that firms choose a value for next period's capital stock that minimizes the following quadratic loss function with capital adjustment costs:

$$\min_{k_{t+1}} \sum_{l=1}^{\infty} \theta^{l} E_t \left[\gamma \left(\Delta k_{i,t+l}\right)^2 + \left(k_{i,t+l} - k_{i,t+l}^{*}\right)^2 \right] \tag{3}$$

[8] This equation can be derived using the standard neoclassical first-order condition for capital for a case where the production function takes the following constant elasticity of substitution (CES) form with Hicks-neutral techological progress:

$$F(A_{i,t}, K_{i,t,}, L_{i,t}) = A_{i,t} \left[\omega_i K_{i,t}^{\frac{\sigma-1}{\sigma}} + (1-\omega_i) L_{i,t}^{\frac{\sigma-1}{\sigma}} \right]^{\frac{\sigma}{\sigma-1}},$$

where σ is the elasticity of input substitution, $K_{i,t}$ is the level of capital, $L_{i,t}$ is the level of the variable input, and $A_{i,t}$ is the level of technology. For simplicity, we suppress all constants and industry fixed effects, and generalize by allowing for a non-unit coefficient on output.

[9] This form of the user cost corresponds to a case where firms are price takers, firms face no explicit internal adjustment costs, and capital requires one period to install before it becomes productive. The present value of future tax shields from capital consumption allowances is allowed to vary by period and by industry.

where $0 \in \theta$ is an appropriate discount rate and where the frictionless stock $k^*_{i,t}$ evolves according to some stochastic process that is known and treated as given by firms.[10] This objective function is the present value of all current and future costs associated with (a) adjusting the capital stock and (b) deviating from the frictionless stock, where the non-negative parameter γ determines the relative importance of these two considerations. The first-order condition for this dynamic optimization problem yields a second-order linear difference equation for capital that can be solved to obtain a non-explosive solution of the form:

$$k_{i,t+1} = \Gamma^B(L)\Gamma^F(L^{-1})k^*_{i,t+1}, \tag{4}$$

where L is the lag operator, $\Gamma^B(L)$ and $\Gamma^F(L^{-1})$ are the following *backward-* and *forward-stable* polynomials in the lag operator:

$$\Gamma^B(L) = \frac{1-\lambda}{1-\lambda L}, \qquad \text{and} \qquad \Gamma^F(L^{-1}) = \frac{1-\theta\lambda}{1-\theta\lambda L^{-1}}, \tag{5}$$

and $\lambda \in (0,1)$ is the stable root. These polynomials satisfy the restriction that $\Gamma^B(1) = \Gamma^F(1) = 1$, ensuring that each industry's capital stock tracks its frictionless stock in the long run. However, in any given period, the optimal level of capital reflects both backward- and forward-looking considerations. To see this, define the *capital target* as

$$k^{**}_{i,t+1} = \Gamma^F(L^{-1})k^*_{i,t+1} = (1-\theta\lambda)\sum_{l=0}^{\infty}(\theta\lambda)^l E_t\left[k^*_{i,t+1+l}\right], \tag{6}$$

so that equation (4) simplifies to the following backward-looking partial-adjustment process:

$$k_{i,t+1} = \Gamma^B(L)k^{**}_{i,t+1}. \tag{7}$$

This shows that the capital stock adjusts in each period towards a target k^{**} that is a weighted average of the firm's expected frictionless capital stocks in all future periods. The weight associated with the anticipated stock at time horizon l is proportional to the factor $(\theta\lambda)^l$, the magnitude of which is negatively related to the time horizon, the time preference parameter θ, and a parameter

[10] This can be thought of as a quadratic approximation to the typical objective function that represents the expected present value of all future net cash flows. For ease of exposition, we assume that the firm only faces costs for changing its level of capital. However, we would obtain a similar with similar long-term properties, but more complicated short-term dynamics, if the objective function was generalized to incorporate investment adjustment costs, i.e. costs of varying the level of investment (or $\Delta^2 k_{i,t}$).

λ that governs the "stickiness" of the capital stock and that has a magnitude that is positively related to the degree of adjustment costs.

Tevlin and Whelan [2003] argue that the forward-looking nature of the capital target is crucial for empirical estimation because the response of capital to an unanticipated change in fundamentals will be closely related to the expected persistence of the disturbance. We incorporate this possibility by letting the evolution of the neoclassical fundamentals in each industry be determined by the following reduced-form univariate processes:

$$C^v(L)v_{i,t} = e^v_{i,t} \qquad \text{for} \qquad v = \{y, u, a\}, \qquad \text{where} \qquad C^v(x) = 1 - \sum_{m=1}^{\infty} c^v_m x^m. \tag{8}$$

Each of the variables e^v is a serially-uncorrelated forecast error that, in general equilibrium, is a linear combination of unobserved structural disturbances. We assume that these three lag polynomials are well-behaved in the sense that, for all v in $\{a, u, y\}$, $C^v(1) \geq 0$. In the Appendix, we show that equations (1), (6) and (8) can be combined to obtain the following expression for the target capital stock as a distributed lag of its fundamentals:

$$k^{**}_{i,t+1} = \beta D^y(L)y_{i,t} - \sigma D^u(L)u_{i,t} + (\sigma - 1) D^a(L)a_{i,t} \tag{9}$$

In this expression, each of the long-run responses $D^v(1)$ reduces to:

$$D^v(1) = \frac{C^v(\theta\lambda) - C^v(1)}{C^v(\theta\lambda)} \qquad \text{for} \qquad v = \{y, u, a\},$$

where, in each case, it can be shown that $0 \leq D^v(1) \leq 1$. The significance of these latter expressions can be illustrated using the following extreme cases, where, without loss of generality, we focus on a disturbance to output. At the one extreme, where all shocks to output are permanent, the polynomial $C^y(L)$ has a unit root so $C^y(1) = 0$. It follows from the formula above that $D^y(1) = 1$, so that the shock has the same long-run effect on both the capital target and the frictionless capital stock. At the other extreme, where all shocks are so transitory that they only affect the level of output for one period, then $C^y(1) = C^y(\theta\lambda) = 1$ and $D^y(1) = 0$. In this case, disturbances to output have no effect on the capital target because they die off before firms have an opportunity to react.[11]

[11] The effect of the shock is literally zero in this case because of the one period time to build requirement. If there were no time to build requirement, so that firms can use capital in the same period that they acquire it, then the shock would generate a non-zero response of a magnitude that is negatively related to the stickiness parameter λ.

Using equations (4) and (9), the capital stock in each industry can be expressed as

$$k_{i,t} = \beta G^y(L)y_{i,t} - \sigma G^u(L)u_{i,t} + (\sigma - 1)\, G^a(L)a_{i,t}, \tag{10}$$

where

$$G^v(x) = \Gamma^B(x)D^v(x), \qquad \text{for} \qquad v = \{y, u, a\}\,.$$

Note that these lag polynomials contain the same polynomials that appear in the expression for the capital target, so the long-run sensitivity of capital to a given change in fundamentals will be closely related to how that change affects the capital target. Since the terms in the backward-looking polynomial $\Gamma^B(L)$ always sum to 1, both the capital target and the optimal capital stock have the same sensitivity to shocks in the long run.

The formulation above suggests a number of important implications for the estimation of capital elasticities. First, the long run response of capital to changes in any of the frictionless fundamentals is closely related to the expected persistence of the innovation. The long run response of capital to a change in any of these fundamentals will only be the same as the frictionless elasticities in equation (1) if the process that describes the evolution of that fundamental contains a unit root. Second, the last term of this equation includes current and lagged values of the unobserved fundamental $a_{i,t}$. Since current and lagged values of this factor are likely to be correlated with the observed investment fundamentals, econometricians should be mindful of the potential for endogeneity when estimating specifications similar to the form in equation (10). In addition, economists often think of technology as an integrated process. If this is true, the cointegration approach, as employed by Caballero [1994] and Schaller [2006], could yield spurious results unless technology shocks do not augment capital.[12] A third implication is the potential for small sample bias in the estimates, which is exascerbated in this setting because the term in 10 involving the unobserved factor $a_{i,t}$ is likely to be serially correlated.

As in Schaller [2006], our estimation strategy aims to identify the long-run elasticity of capital to the user cost by choosing a sample in which the evolution of the user cost is more likely to be

[12] Another apparent possibility is to assume, as is implicit in Caballero [1994] and Schaller [2006], that the production function is Cobb-Douglas. However, this is not a real possibility, because it implicitly imposes the restriction that σ is 1, so that the elasticity is no longer an undetermined parameter. Any other finding would either contradict the assumption that the production function is Cobb Douglas or be inconsistent with the presumption that the user cost follows an integrated process–a necessary condition for the cointegrating regression to be valid.

determined by factors outside of the domestic economy. We think that it is plausible that forecast errors for the user cost ($\varepsilon_{i,t}^{u}$) are orthogonal to current and lagged changes in the unobserved productivity factor $a_{i,t}$ because South Africa is a price taker in the international markets.[13] Even if this presumption is true, there remains a possibility that there are some productivity shocks that increase capital demand in all countries, and are therefore correlated to movements in the various factors that compose the user cost of capital. This possibility should be limited to some extent in the South African case because its economy is relatively isolated from the world's largest economies by geography. We can also at least partially account for this problem by estimating our regression using a panel of industries, where it is more likely that the disturbances stem from idiosyncratic factors, and by controlling for aggregate shocks that can be identified in the cross-section dimension.

3 Estimation and Results

3.1 Specification and Data

Our primary procedure is to estimate a regression specification of the basic form:

$$\Delta k_{i,t} = \eta_i + d_t + N^y(L)\Delta y_{i,t} + N^u(L)\Delta u_{i,t} + \epsilon_{i,t} \tag{11}$$

for industries $i = 1, ..., N$, where η_i is an industry fixed effect, $\epsilon_{i,t}$ is an unexplained residual, and d_t is an embargo dummy for the period from 1985:Q3 to 1994:Q2.[14] The form of this equation can be justified by taking the first difference of equation (10), where we let $\eta_i + \epsilon_{it}$ denote the first difference of the unobserved term $(\sigma - 1)\, G^a(L)a_{i,t}$ and noting that the lag polynomials take the structural form:

$$N^y(L) = \beta G^y(L), \qquad \text{and} \qquad N^u(L) = -\sigma G^u(L).$$

In order to estimate this equation, we assume that these lag polynomials are of finite order, so that the long run response of capital to each fundamental can be estimated by summing the estimated

[13] In fact, we think that this condition is much more likely to be true for South Africa than for Schaller's country of analysis, Canada. While it is at least somewhat plausible that Canada–a G7 member–treats prices of financial capital and goods as given, it is much less likely that these prices are exogenous due to Canada's close economic ties to the United States, the world's largest economy.

[14] We interpret the beginning of the embargo as September 1985, when official sanctions against South Africa were enacted by the European Community and the United States. The end of our embargo period is the quarter in which the first all-race democratic elections were held in South Africa. Shortly thereafter, the United Nations adopted a resolution for all of its members to end economic sanctions against the country.

parameters of the corresponding distributed lag function.[15] Under the maintained assumption that innovations to the user cost are exogenous in the non-embargo period, then the parameters in the lag polynomial $N^u(L)$ are identified and can be estimated consistently using standard least-squares techniques. This remains true even if the parameters of the polynomial $N^y(L)$ are not identified, which is likely to be the case throughout the sample period because of the apparent correlation between the structural residual ϵ_{it} and the current and lagged values of output.

For all of our regressions, we use a quarterly panel of twenty-four manufacturing industries at the two-digit level that extends from 1970:Q1 to 2001:Q4.[16] Industry-level estimates of the capital stock, fixed investment, consumption of fixed capital, price deflators, and output were obtained from South Africa Trade and Industrial Policy Strategies (TIPS). Quarterly data on interest rates, corporate tax rates, tax credits, capital depreciation allowances, and the aggregate price level are from the South African Reserve Bank. We calculate the user cost of capital in each period and for each industry using equation (2), the components of which were determined as follows. The real borrowing cost (r_t) in each quarter is the end-of-quarter nominal prime overdraft rate (a short-term rate charged to commercial banks), to which we add a fixed risk premium and then deduct expecated inflation measured by the realized growth rate in the GDP price deflator over the coming four quarters.[17] The price of investment goods ($\overline{p}_t^k$) is the industry's price deflator net of our estimate of the present value of the tax shields from depreciation allowances.[18] The depreciation rate for each industry in each quarter ($\delta_{i,t}$) is obtained by dividing the industry's

[15] Adding autoregressive terms had very little effect on our long-run elasticity estimates, once we had included the large number of lags of output and the user cost that are in our baseline specification.

[16] These twenty-four industries are: (1) food, (2) beverages, (3) textiles, (4) wearing apparel, (5) footwear, (6) wood and wood products, (7) paper and paper products, (8) printing, publishing and recorded media, (9) coke and refined petroleum products, (10) basic chemicals, (11) other chemicals and man-made fibers, (12) rubber products, (13) plastic products, (14) non-metallic minerals, (15) basic iron and steel, (16) basic non-ferrous metals, (17) metal products excluding machinery, (18) machinery and equipment, (19) professional and scientific equipment, (20) motor vehicles, parts and accessories, (21) other transport equipment, (22) furniture, (23) Electrical machinery and apparatus, and (24) other manufacturing. We excluded four industries (tobacco, leather and leather products, glass and glass products, television, radio and communication equipment, and electrical machinery and apparatus) because their data were either of poor quality or did not exist for fixed investment.

[17] We use the four-quarter rate of change, rather than the one-quarter change that matches the frequency of our sample, in order to state capital gains at an annual frequency and to minimize variations owing to price seasonalities.

[18] More specifically, $\overline{p}_{i,t} = p_{i,t}\,(1 - \tau_t z_{i,t})$, where $z_{i,t}$ is the present value of all the future depreciation allowances associated with a unit of investment at time t. This factor was approximated using the formula $z_t = \frac{\delta_{i,t}}{i_t + \delta_{i,t}} = \sum_{j=1}^{\infty} \delta_{i,t}\,(1+i_t)^{-j}\,(1-\delta_{i,t})^{j-1}$, where i_t is the nominal interest rate. This formula implicitly assumes that firms expect interest rates, the corporate tax rate, and the rate of depreciation to remain constant at their current levels.

consumption of fixed capital by its capital stock at the end of the previous quarter. Finally, we proxy for the anticipated rate of appreciation in investment goods using the realized rate of appreciation recorded over the following four quarters. Our estimates of the user cost of capital in each quarter differ across industries because of differences in the relative price of capital and the rate of depreciation.

It is worth noting that the specification shown above does not include any lagged values of the dependent variable as regressors, unlike the specifications used by Tevlin and Whelan [2003] and others. In principle, the two approaches should be roughly equivalent, since one can recover an autoregressive specification by a simple rearrangement of our structural equation (10).[19] Our approach allows the dynamic response of capital to changes in fundamentals to take a general form that, unlike the autoregressive specification, does not impose a geometric rate of decay. In addition, autoregressive specifications may be disadvantageous if the structural error is serially correlated, because the estimated autoregressive parameters would be inconsistent. The autoregressive formulation may also be prone to errors-in-variables problems that could arise if capital growth is measured with error. Measurement error in the rate of capital growth should be relatively innocuous in our specification because it is absorbed by the regression residual. That said, our methodology has the disadvantage that it may miss any portion of the capital response because we can only include a finite number of lags in our regression. In order to limit this possibility, we include an unusually large number of lags in our regression specifications.[20]

For the reasons explained in the previous section, the long-run elasticities estimated using equation (11) can only be interpreted as the frictionless elasticities in equation (1) when the fundamental in question follows a unit root process. We consider the unit root issue in Table 1, which shows results of panel unit root tests for each of our variables of interest: capital, the user cost, and output. The table shows results from a number of alternative test procedures that differ in formulation, robustness, and the maintained null hypothesis. Four of these tests maintain a null that there is a unit root in the series of interest for all industries in the panel: the Levin, Lin and Chu [2002], Im, Pesaran and Shin [2003], Maddala and Wu [1999] and Pesaran [2003] tests. The Im, Pesaran and Shin and Pesaran versions are the most robust of these tests because they allow the possibility

[19]For instance, one could obtain an autoregressive form for capital by multiplying equation 10 by the inverse of the backward-looking polynomial $\Gamma^B(L)$.

[20]When we increased the number of included lags in these regressions it had little effect on our estimates, and the direction of the effect depended on the specification.

that the residuals are correlated in the cross-sectional dimension; the Pesaran test allows this cross-sectional correlation to take a less restricted form.[21] The fifth test, from Hadri, maintains the null that the variable does not follow a unit root process in any of the industries of the panel. The results for almost all of these tests suggest that the processes for both output and capital contain unit roots. In contrast, the panel tests for the user cost of capital are quite mixed. Nonetheless, the results of the most robust panel tests seem to offer some limited support for the existence of a unit root in this process. Therefore, even though the evidence for a unit root in the user cost is not overwhelming, these results offer at least some support for interpreting our estimate of the long-run user cost elasticity of capital as an estimate of the frictionless elasticity parameter σ.

3.2 Estimates Using the Entire Data Sample

To fix ideas, we begin by reporting results from regression specifications that do not explicitly account for the heightened potential of user cost endogeneity during the embargo portion of the sample period. These results serve as a basis of comparison both to previous studies and for the results we report later that account for the apparent endogeneity of the user cost during the embargo.

Table 2 shows our regression estimates using the difference specification described above. All of these specifications include contemporanous values of the user cost and output, along with 32 lags of each of these variables.[22] The regressions also include an embargo fixed effect and a constant (not shown). The results shown in columns (1) to (4) of the table are estimated using pooled OLS, while the results reported in columns (5) to (8) control for industry fixed effects. The standard errors reported for each estimate are robust to the possibility of correlation between the residuals in the cross-sectional dimension. All of these estimates include in the user cost our proxy of anticipated capital gains, but including this term had only a small effect on our estimates.[23]

[21] Specifically, the Im-Pesaran-Shin test controls only for a fixed effect for each time period that is common across groups, while the Pesaran test allows for a more general form of dependence.

[22] We chose a lag length sufficient to capture all statistically significant coefficients on lags of the user cost and output.

[23] It is not obvious in practice whether it is better to leave the anticipated capital gains term out of the user cost or to proxy for it using realized capital gains, since both approaches may introduce some measurement error. However, we think that the proxy approach is more appropriate when there are persistent movements in the relative price of business investment goods. It is well known that there has been a noticable downward trend in the relative price of business investment goods in the U.S. aggregate data from the early 1980s onward. This downward trend is also evident in the South African data from the manufacturing sector, although it is temporarily reversed during the

Column (1) shows regression results from our baseline OLS specification, while column (5) shows estimates from the same specification with fixed industry effects. Our estimate of the long run user cost elasticity of capital for the baseline case is about -0.66, and is statistically different from zero at significance levels well below one percent. Including industry fixed effects diminishes the absolute magnitude of the user cost elasticity estimate only slightly. The estimates of the long run output elasticity in these two specifications–which are probably inconsistent for the reasons discussed above–are 0.67 for the baseline specification and 0.59 with industry fixed effects. Both of these estimates are statistically distinguishable from zero at standard significance levels, but tests also show that they are well below 1.0 (the neoclassical benchmark) at low levels of significance.

Figure 2 shows the point estimates of the response of capital growth to the user cost at each lag and 95 percent confidence intervals of these estimates, using the OLS benchmark specification reported in column (1). This impulse response function shows a very distinct hump-shape that reaches a rough plateau between the 8th and 17th quarters, and, with the exception of the first and last few lags, the individual responses are very significant. Figure 3 plots these same set of estimates restated as a cumulative response, showing that the total response is quite drawn out. One particularly notable aspect of these estimates is that the marginal response function is not diminishing and concave, as one would expect with convex costs of adjusting the level of capital. Among other things, this characteristic would seem consistent with adjustment costs associated with altering capital growth (the investment rate).

The estimates in the remaining columns of Table 2 use some additional methods to control for endogeneity. Columns (2) and (6) restrict the baseline specification so that the long-run output elasticity is 1.0, as is the case of constant returns to scale. These estimates should reveal whether the endogeneity of the output term affects the user cost elasticity, as one might expect if our identifying assumption was not valid so that the true residual $\epsilon_{i,t}$ is correlated to both output and the user cost. The elasticity estimates from these two specifications are a little larger in absolute magnitude than in the baseline specification. This suggests that, if our estimates from the baseline specification are inconsistent, they may be too small. However, the point estimates from these restricted regressions are well within any standard confidence interval of the baseline estimates, so the difference is probably not statistically meaningful. Columns (3) and (7) show estimates that include quarterly dummies that are common across industries, and should therefore control

embargo period.

for any remaining "aggregate" component of the true residual that affects capital accumulation in all industries.[24] The point estimates of the user cost elasticity obtained using this specification are somewhat smaller in absolute magnitude than those from the baseline specification, but are well within a standard confidence interval of the baseline estimates. The final set of estimates reported in columns (4) and (8) include both the quarterly dummies and the restriction on the long-run output elasticity. These estimates are roughly in line with the baseline cases.

To our knowledge, these user cost elasticity estimates are at the high end of those seen in the literature, especially when one takes into account that they include nonresidential structures in the measure of capital. For instance, the cointegration methodology employed by Caballero [1994], using aggregate U.S. data, and Schaller [2006], using aggregate Canadian data, yield estimates of the user cost elasticity that approach or exceed 1.0 in absolute value, but neither of these estimates include nonresidential structures in the measure of capital. When these authors use a measure of capital that includes nonresidential structures, they find that the user cost elasticity is essentially zero. Using a similar specification to our own that is estimated using data from a huge panel of U.S. firms, Chirinko, Fazzari and Meyer [1999] obtain some estimates that are similar in magnitude and precision to our estimates, even though their "preferred" estimates (obtained using an instrumental variables technique) range between -0.5 and 0. However, estimates of this magnitude are far from representative. Usually, empirical estimates of the user cost elasticity are close to zero or have the wrong sign, and are seldom statistically significant.

3.3 Estimates Using a Panel Cointegration Approach

An alternative specification for estimating the long-run elasticities of capital to the user cost and output is to amend equation (10) slightly into a cointegrating specification of the form

$$k_{i,t} = \beta y_{i,t} - \sigma u_{i,t} + e_{i,t}, \tag{12}$$

where

$$e_{it} = \beta \left[G^y(L) - 1 \right] y_{it} - \sigma \left[G^u(L) - 1 \right] u_{it} + (\sigma - 1)\, G^a(L) a_{i,t}.$$

[24] In all of the specifications that include the quarterly dummies, we leave out the embargo fixed effect because it cannot be separately identified.

This is a valid cointegrating specification provided that capital is integrated and the reduced-form error term $e_{i,t}$ is not integrated–a condition that requires both output and the user cost to follow unit root processes and that the unobserved technology term $a_{i,t}$ is either stationary or does not really exist.[25]

Assuming that the cointegrating specification is valid, large-sample estimates using this approach should be superior to those from our difference specification. This is because of the well-known property that the structural parameters from a cointegrating regression are super-consistent, even if the presence of endogeneity between the regressors and the residual term. So, in a sufficiently large sample, this method should recover the true structural parameters even if our assumption about the exogeneity of the user cost during the non-embargo period does not hold true. That said, the structural form of our residual suggests that the estimates from this regression may suffer from small sample bias.[26] We correct for this possibility by estimating our cointegrating regression by dynamic panel OLS with a large number of leads and lags of the first-difference of the independent variables.[27]

In order to determine whether the cointegrating specification is valid, we test whether the fitted residual terms e_{it} from estimates of equation (12) are stationary using the same set of tests used in Table 1. As discussed earlier, the panel unit root tests shown in Table 1 provided strong support that there is a unit root in the processes for output and capital, and some partial support for a unit root in the process for the user cost of the frictionless capital. For robustness in our cointegration tests, we used two separate specifications: The first specification treats the long-run output elasticity β as a free variable and estimates it in the regression, while the second specification

[25] These conditions ensure that $e_{i,t}$ is stationary as follows. As shown earlier, both $G^y(1) = 1$ and $G^y(1) = 1$ provided that the autoregressive polynomials for these variables shown in equation (8) contain unit roots. This implies that the lag polynomials $[G^y(L) - 1]$ and $[G^u(L) - 1]$ that multiply output and the user cost above must contain unit roots, which ensures that the first two components of $e_{i,t}$ are stationary. The lack of a unit root in $a_{i,t}$ ensures that the final component of $e_{i,t}$ is stationary.

[26] Caballero [1994] shows simulation results that suggest that the degree of small-sample bias can be considerable in the univariate case. We repeated these simulation experiments in a panel context (not shown) and came to a similar conclusion.

[27] Kao and Chiang [2000] show that estimates of the coefficients in a cointegrating regression from the panel OLS estimator have a biased asymptotic distribution. Simulations in their paper show that the dynamic panel OLS (DOLS) estimator has only a small bias for samples with cross-section and time dimensions similar to our panel, and that this estimator outperforms alternative estimators like pooled OLS and fully-modified OLS. Their simulations also show that conventional standard error estimates from the DOLS regression have only a small bias, in sharp contrast to these alternative estimators.

is constrained so that this long-run output elasticity is one (as in Caballero [1994] and Schaller [2006]). The results of these tests are reported in Table 3. As one can see, a handful of the tests do support the null of cointegration, especially for the specifications that restrict the magnitude of the long run output elasticity. However, the large majority of these tests, including the most robust version of the test from Pesaran [2003], fail to reject the null of no cointegration. Although these results are not conclusive, the weight of the evidence, taken together with the results of the unit root tests described earlier, suggests that the cointegrating regression approach may not be valid. These findings are also consistent with our priors about the implausibility of a stationary technology factor $a_{i,t}$. That said, our panel tests may still lack the power needed to reject the null of no cointegration, so we proceed with this approach as a robustness check for our preferred difference specification.

Our estimates using this cointegrating specification are reported in Table 4. Columns (1) and (2) of the table report estimates for a specifications that restrict the long-run elasticity of capital to output to be 1. However, the specification in column (1) includes only leads and lags of the change in the user cost as dynamic OLS terms, while column (2) also includes leads and lags of the change in output. The estimates of the user cost elasticity from these specifications are both in the vicinity of -1.0, and are highly significant. The final specification—shown in column (3)—estimates the long run output elasticity of capital as a free variable, and also includes leads and lags of the change in both output and the user cost as dynamic OLS terms. The estimated elasticities from this specification are also highly significant and are somewhat larger in absolute magnitude than our estimates from the difference specification. The user cost elasticity from this specification (-0.94) is a little smaller in absolute magnitude than the constrained estimates in columns (1) and (2) , but is also not statistically distinguishable from -1.0.

3.4 Estimates that Control for the Embargo Period

We now move on to estimates that account for the possibility that the user cost of capital is subject to a heightened degree of endogeneity during the embargo period. To account for this possibility, we augment our original regression specification to include terms that interact the observable explanatory variables with our embargo dummy. Our formulation for this regression is:

$$\begin{aligned}\Delta k_{i,t} &= d_t + N^y(L)\Delta y_{i,t} + M^y(L)\left(d_t\Delta y_{i,t}\right) \\ &\quad + N^u(L)\Delta u_{i,t} + M^u(L)(d_t\Delta u_{i,t}) + \epsilon_{i,t}, \end{aligned} \tag{13}$$

so that the embargo affects the entire long run relationship between capital and its fundamentals, but only for observations of these fundamentals that occur within the embargo period. As in our previous difference specification, we estimate this regression using the contemporaneous observations and 32 lags of each the fundamentals (including the interactions with the embargo dummy). Given these estimates, we determine the long-run elasticity of capital with respect to the user cost and output by calculating $N^y(1)$ and $N^u(1)$, respectively. The marginal effect of the embargo-period data on our long-run user cost and output elasticities are calculated as $M^y(1)$ and $M^u(1)$, respectively, while the long-run elasticities to these fundamentals during the embargo-period are $N^y(1) + M^y(1)$ and $N^u(1) + M^u(1)$. Importantly, we do not interpret any discrepancy in these embargo and non-embargo estimates to changes in the *structural* relationship between capital and its observable fundamentals shown in equation (10). Rather, these discrepancies owe only to the change in the user cost from an exogeneous to an endogenous variable during the embargo period, so that movements in the user cost becomes correlated to the residual term in equation (11). Our assumption that the embargo increased the endogeneity of the user cost can be verified by testing that $M^u(1) < 0$. We can also test whether the long-run elasticity that we estimate for the embargo period is significantly different from zero by looking at the sum $N^u(1) + M^u(1)$.

Results using this specification are shown in Table 5. The results reported in this table are analogous to those reported in Table 2, with the first four columns showing estimates obtained using a pooled OLS specification, and the second four columns showing estimates that control for industry fixed effects. As in Table 2, columns (1) and (5) use a baseline specification that includes only an embargo fixed effect, while columns (2) to (4) and (6) to (8) include the various controls for endogeneity described earlier. For each of these specifications, the estimate of the long-run user cost during the non-embargo period is larger in absolute magnitude than the corresponding estimates in Table 2 that do not account for embargo interaction effects. All of these estimates are highly significant, and range in magnitude from around -0.71 for the baseline cases in columns (1) and (5) to as high as -0.85 for the estimates in columns (4) and (8) that include quarterly dummies and impose a unit restriction on the long-run output elasticity. In all of our specifications, we found that the embargo period data caused a large and statistically significant drop in the absolute magnitude of the estimated user cost elasticity. These drops were so considerable that the estimated elasticities during the embargo period are quite small, ranging between around -0.05 in column (3) to -0.25 in column (4). In most of these cases, the elasticity during the embargo period is not statistically distinguishable from zero, although the user-cost elasticity is still significant at around five percent or lower in columns (4) and (8)–which control for quarterly dummies and impose the unit restriction on the output elasticity. The table also provides some evidence that the embargo-period data may attenuate our estimate of the output elasticity of capital demand, albeit to a lesser extent than the user cost elasticity. Yet even after accounting for this attenuation, movements in output still had a statistically significant long-run effect on capital accumulation during the embargo period.

These results are consistent with our intuition that the user cost became more endogenous during the embargo period, heightening the degree of inconsistency in our estimates of the user cost elasticity over that period. These findings may help explain why most previous empirical studies of investment using U.S. data tend to find very little role for the user cost in determining the size of the capital stock.[28]

4 Conclusion

In a closed economy or in a large open economy, both the capital stock and the user cost of capital are jointly determined by domestic demand and supply equilibrium that equates the marginal product and marginal opportunity cost of capital services. This simultaneity introduces inconsistency into estimates of the user cost elasticity. Our study expands on an insight by Schaller [2006] that in a small open economy where the price of investment goods and the interest rate are largely determined by foreign markets, the user cost is less likely to be endogenous, thereby allowing for consistent estimates of the user cost elasticity. Using a quarterly panel of two-digit manufacturing data from South Africa, we obtain estimates of the user cost elasticity that are large and reasonably close to the Cobb-Douglas benchmark. This study is the first to document such a large user cost elasticity for a broad measure of business capital that includes both equipment and structures. Unlike previous studies that obtained large elasticities for just equipment capital, our study does

[28] For instance, see the review by Chirinko [1993].

not rely on the existence of a cointegrating relationship between capital, output, and the user cost.

The economic embargo that the world imposed on South Africa between 1985 and early 1994, which forced the economy to behave much more like a closed economy provides another test of Schaller's hypothesis. We find evidence that during the embargo period, the magnitude of the estimated user cost elasticity fell considerably—to magnitudes that are consistent with previous studies in which the user cost was likely endogenous. This finding underscores the importance of identification when forming estimates of the user cost elasticity.

A Appendix: Omitted Proof

Obtaining the expression in equation (9) boils down to down to calculating:

$$(1-\theta\lambda)\sum_{l=0}^{\infty}(\theta\lambda)^{l}E_{t}\left[v_{i,t+1+l}\right]$$

for each v in $\{y,u,a\}$, where the evolution of these variables are described by the univariate process shown in equation (8). As a first step, note that, by the law of iterated expectations

$$(1-\theta\lambda)\sum_{l=0}^{\infty}(\theta\lambda)^{l}E_{t}\left[v_{i,t+1+l}\right]=(1-\theta\lambda)\,E_{t}\left[\sum_{l=0}^{\infty}(\theta\lambda)^{l}E_{t+1}\left(v_{i,t+1+l}\right)\right], \tag{A1}$$

Hansen and Sargent [1980] show that the solution for the bracketed term on the right-hand side of this equality is

$$\frac{1}{1-C^{v}(\theta\lambda)}\left[1+\sum_{j=1}^{\infty}\left(\sum_{k=j+1}^{\infty}(\theta\lambda)^{k-j}\,c_{k}^{v}\right)L^{j}\right]v_{t+1}. \tag{A2}$$

Substituting this solution into equation (A1) and distributing through the expectations operator E_t:

$$(1-\theta\lambda)\sum_{l=0}^{\infty}(\theta\lambda)^{l}E_{t}\left[v_{i,t+1+l}\right]=\frac{1-\theta\lambda}{1-C^{v}(\theta\lambda)}\left[E_{t}v_{t+1}+\sum_{j=1}^{\infty}\left(\sum_{k=j+1}^{\infty}(\theta\lambda)^{k-j}\,c_{k}^{v}\right)L^{j}v_{t+1}\right] \tag{A3}$$

Now use equation (8) to substitute out the conditional expectation $E_t v_{t+1}$ and simplify to obtain

$$(1-\theta\lambda)\sum_{l=0}^{\infty}(\theta\lambda)^{l}E_{t}\left[v_{i,t+1+l}\right]=\frac{1-\theta\lambda}{1-C^{v}(\theta\lambda)}\left[\sum_{j=1}^{\infty}c_{j}^{v}L^{j-1}+\sum_{j=1}^{\infty}\left(\sum_{k=j+1}^{\infty}(\theta\lambda)^{k-j}\,c_{k}^{v}\right)L^{j-1}\right]v_{t}, \tag{A4}$$

Equation (9) follows from the equation above by defining:

$$D^{v}(L)\equiv\left[\sum_{j=1}^{\infty}c_{j}^{v}L^{j-1}+\sum_{j=1}^{\infty}\left(\sum_{k=j+1}^{\infty}(\theta\lambda)^{k-j}\,c_{k}^{v}\right)L^{j-1}\right].$$

To obtain the expressions for the long run response $D^v(1)$, note that, by the equation above,

$$D^{v}(1)=\frac{1-\theta\lambda}{C^{v}(\theta\lambda)}\left[1-C^{v}\left(1\right)+\sum_{j=1}^{\infty}\left(\sum_{k=j+1}^{\infty}(\theta\lambda)^{k-j}\,c_{k}^{v}\right)\right]. \tag{A5}$$

Rearranging terms in this equation, it can be shown that

$$D^v(1) = \frac{1-\theta\lambda}{C^v(\theta\lambda)}\left[1 - C^v(1) + \sum_{j=1}^{\infty}(\theta\lambda)^j \sum_{k=j+1}^{\infty} c_k^v\right] \tag{A6}$$

which, once again, can be simplified further using equation (8) to yield

$$D^v(1) = \frac{1-\theta\lambda}{C^v(\theta\lambda)}\left[1 - C^v(1) + \sum_{j=1}^{\infty}(\theta\lambda)^j \left(1 - C^v(1) - \sum_{r=1}^{j} c_r^v\right)\right]. \tag{A7}$$

As a final step, note that:

$$\sum_{j=1}^{\infty}(\theta\lambda)^j = \frac{\theta\lambda}{1-\theta\lambda},$$

and that, by expanding and collecting terms:

$$(1-\theta\lambda)\sum_{j=1}^{\infty}(\theta\lambda)^j \sum_{r=1}^{j} c_r^v = 1 - C^v(L).$$

Using these two expressions, one can show that equation (A7) is equivalent to the expression for $D^v(1)$ in equation (9).

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Figure 1: Current Account Balance to GDP Ratio

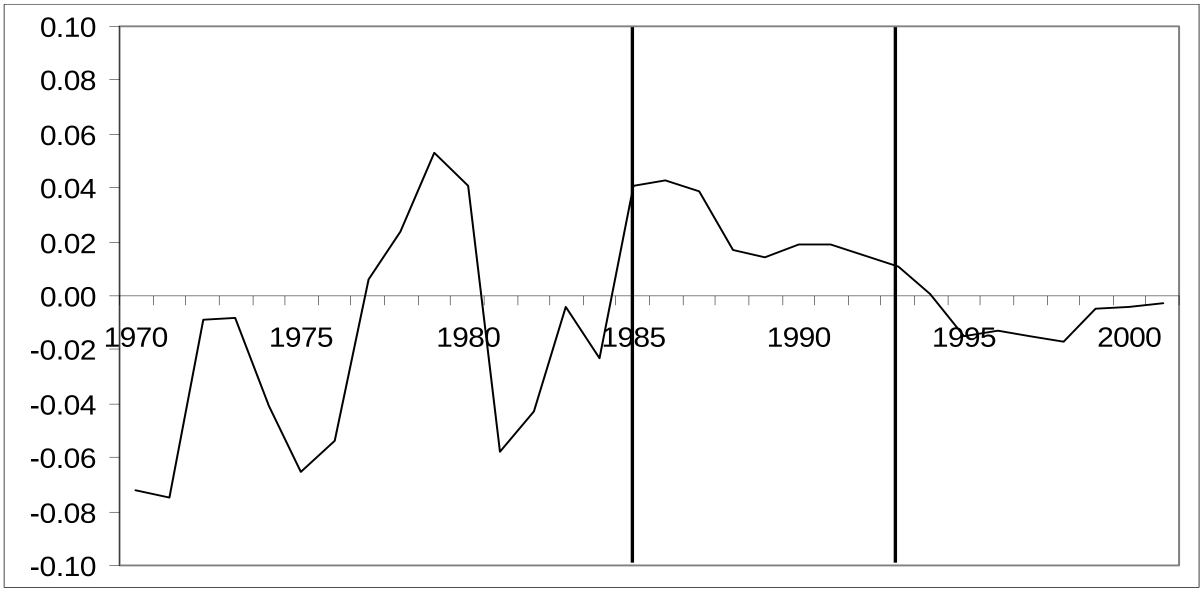

Figure 2: Estimated Marginal Response of Capital to User Cost with 95 percent Confidence Interval

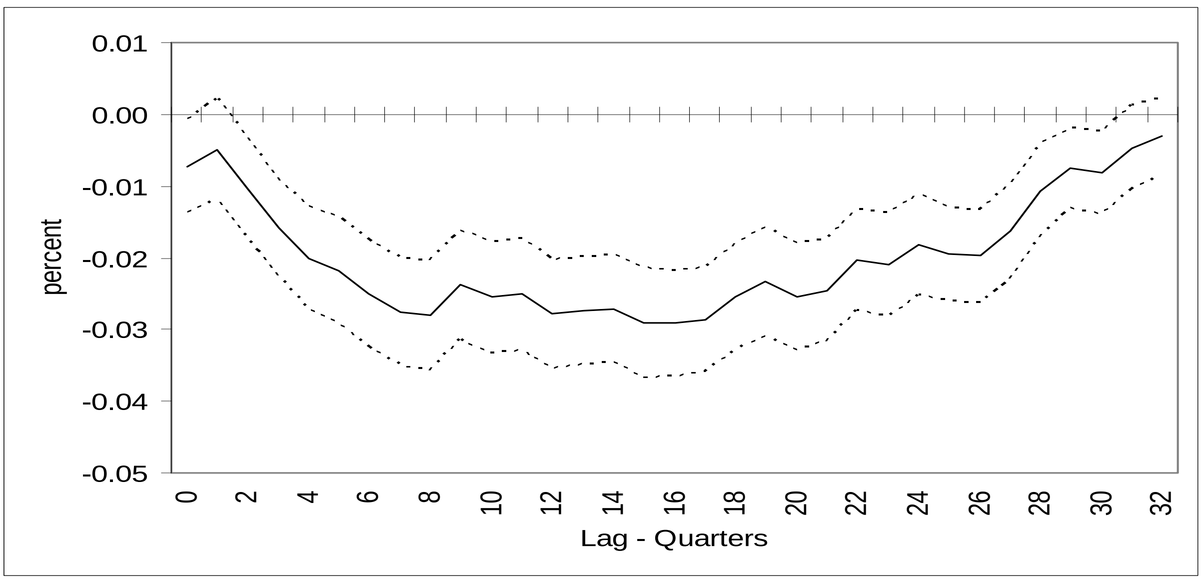

Figure 3: Estimated Cumulative Marginal Response of Capital to User Cost with 95 percent Confidence Interval

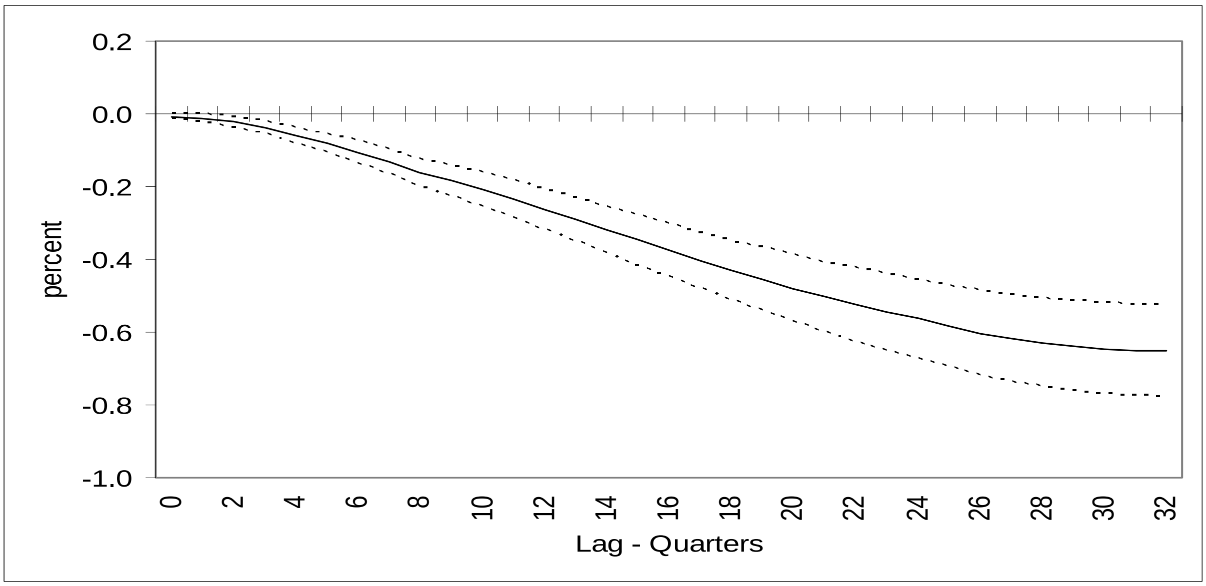

Table 1: Panel Unit Root Tests (p-value for H_0 with constant term and no time trend)*

Test	H_0	User Cost	Output	Capital	$\frac{\text{Capital}}{\text{Output}}$
Levin-Lin-Chu	common unit root	0.000	0.000	0.100	0.136
Im-Pesaran-Shin	common unit root	0.003	0.772	0.963	0.448
Maddala-Wu	common unit root	0.002	0704	0.945	0.048
Pesaran	common unit root	0.289	0.998	0.997	0.437
Hadri	no unit roots for any group	0.000	0.000	0.000	0.000

All variables are in logs. In all cases, a preliminary regression was run to partial out group effects and a fixed that is common across groups. Each test is augmented to account for serial dependence in the fitted error. The Im-Pesaran-Shin test is robust for cross-sectional of errors owing to common fixed time effects, while the Pesaran test is robust for more generic cross-sectional error correlation.

*We obtain similar results when the tests include both a constant and time trend.

Table 2: Results from OLS Difference Specification using Panel Data (robust standard error)

	OLS				FE			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Long-Run Output	0.672	1.000	0.683	1.000	0.589	1.000	0.530	1.000
Elasticity	(0.078)	(—)	(0.096)	(—)	(0.089)	(—)	(0.128)	(—)
Long-Run User Cost	-0.655	-0.714	-0.581	-0.672	-0.635	-0.700	-0.524	-0.647
Elasticity	(0.064)	(0.064)	(0.076)	(0.072)	(0.064)	(0.064)	(0.079)	(0.071)
Embargo	-0.005	-0.003			-0.006	-0.003		
Dummy	(0.001)	(0.004)			(0.001)	(0.001)		
R^2	0.185	—	0.223	—	0.227	—	0.264	—
N	1,963	1,963	1,963	1,963	1,963	1,963	1,963	1,963

All regressions include the contemporaneous observation of the dependent variables and 32 lags. The robust standard errors account for both cross-sectional correlation and heteroskedasticity in the residuals. Results labeled (1) and (5) are for a baseline specification that includes an embargo dummy, while specifications (2) and (6) restrict this baseline specification so that the long-run output elasticity is 1. Specifications (3) and (7) include quarterly dummies in the baseline specification. Specifications (4) and (8) include both restrictions on the long-run output elasticity and quarterly dummies.

Table 3: Panel Cointegration Tests (p-value under H_0)

Test	H_0	$e_{i,t} = k_{i,t} - by_{i,t} - cu_{i,t}$	$e_{i,t} = k_{i,t} - y_{i,t} - cu_{i,t}$
Levin-Lin-Chu	no cointegration	0.150	0.000
Im-Pesaran-Shin	no cointegration	0.559	0.092
Maddala-Wu	no cointegration	0.000	0.000
Pesaran	no cointegration	0.956	0.806
Hadri	cointegration	0.000	0.000

All tests are unit root tests using the fitted residuals from a cointegrating vector estimated by DOLS using 25 leads and lags of the dependent variables. A preliminary regression was run to partial out group means and a fixed embargo effect. The critical values for each test do not account for the generated regressors in the first and second stages described above. Each test is augmented to account for serial dependence. The Im-Pesaran-Shin test is robust for cross-sectional correlation of errors owing to common fixed time effects, while the Pesaran test is robust to more generic cross-sectional error correlation.

Table 4: Panel Cointegration Regression (standard errors)

	(1)	(2)	(3)
Long-Run	1.000	1.000	0.874
Output Elasticity	(—)	(—)	(0.043)
Long-Run	-1.098	-1.017	-0.943
User Cost Elasticity	(0.065)	(0.059)	(0.064)
R^2	0.277	0.538	0.648
N	1,469	1,469	1,469

All specifications are estimated by dynamic pooled OLS with fixed effects. Specifications (1) and (2) are constrained so that the coefficient on output is one, while specification (3) allows the output elasticity to be a free variable. All specifications include as dynamic OLS terms 25 leads and lags of the difference in the user cost. Specification (2) and (3) also include as dynamic OLS terms 25 leads and lags of the difference in output. When we exclude the embargo dummy, the estimates are lower.

Table 5: Results from OLS Difference Specification using Panel Data and Embargo Interactions (robust standard error)

	OLS				FE			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Long-Run Output Elasticity								
Non-Embargo	0.682	1.000	0.751	1.000	0.545	1.000	0.509	1.000
	(0.088)	(—)	(0.107)	(—)	(0.095)	(—)	(0.139)	(—)
Embargo Effect	-0.349	—	-0.429	—	-0.261	—	-0.225	—
	(0.170)	(—)	(0.178)	(—)	0.204	(—)	(0.211)	—
Embargo	0.315	1.000	0.3224	1.000	0.284	1.000	0.284	1.000
	(0.148)	(—)	(0.147)	(—)	(0.187)	(—)	(0.193)	(—)
Long-Run User Cost Elasticity								
Non-Embargo	-0.750	-0.793	-0.810	-0.857	-0.707	-0.767	-0.735	-0.840
	(0.071)	(0.069)	(0.085)	(0.078)	(0.074)	(0.072)	(0.092)	(0.084)
Embargo Effect	0.665	0.512	0.761	0.604	0.627	0.515	0.751	0.624
	(0.089)	(0.083)	(0.134)	(0.126)	(0.089)	(0.085)	(0.145)	(0.140)
Embargo	-0.100	-0.281	-0.048	-0.252	-0.081	-0.154	0.013	-0.132
	(0.097)	(0.095)	(0.122)	(0.116)	(0.095)	(0.099)	(0.130)	(0.123)
Embargo	-0.000	0.000			-0.001	-0.000		
Dummy	(0.002)	(0.001)			(0.002)	(0.001)		
R^2	0.242	—	0.278	—	0.278	—	0.314	—
N	1,963	1,963	1,963	1,963	1,963	1,963	1,963	1,963

All regressions include the contemporaneous observation of the dependent variables and 32 lags. The robust standard errors account for both cross-sectional correlation and heteroskedasticity in the residuals. Results labeled (1) and (5) are for a baseline specification that includes an embargo dummy, while specifications (2) and (6) restrict this baseline specification so that the long-run output elasticity is 1. Specifications (3) and (7) include quarterly dummies in the baseline specification. Specifications (4) and (8) include both restriction on the long-run output elasticity and quarterly dummies.