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TESTING THE STABILITY OF A PRODUCTION FUNCTION WITH URBANIZATION AS A SHIFT FACTOR

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I. INTRODUCTION

Urban economists have long been attempting to answer the question of why cities exist. In particular, recent research has focused on the idea that cities capturing some sort of agglomeration economies-production in urban areas benefit from some increasing returns to scale which are not present in rural environments. Urban economists have gone further to separate these agglomerations into two categories: localization economies (economies of scale present because of industrial clustering in cities) and urbanization economies (economies of scale present because of the overall size of city). Localization economies are external to the firm but internal to the industry. Most of the articles presented with regards to output and urbanization have focused on urbanization economies and not localization economies.

Given that urban structure is hypothesized to influence output levels another related question is how urbanization affects economic development: what is the role of urbanization levels in developing and developed countries? A long-running debate in development economics has been whether developing countries have become over-urbanized. In Moomaw and Shatter (1993) the debate is characterized as that between the traditionalists as represented by Todaro's works, for example his piece (1995) which claims that less-developed countries are over-urbanized, and the modernists as represented in the work of Wheaton and Shishido (1981), which claims that large cities are necessary to realize economies of scale. Unfortunately, there are not many empirical studies which look at this question. Exceptions are Moomaw and Shatter (1993, 1996). Further, the empirical research in the literature is either based on cross-section studies or very limited panels which are unable to truly capture the dynamic nature of the question.

The main idea of the dynamic model we use in this paper comes from the

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neoclassical growth literature which sets out a dynamic optimization problem for the economic individual constrained by some production function. An example of this theory is given by Lucas (1988) and his attempts to explain economic growth in developed and developing countries. His paper is not among those concerned with the convergence theories of growth such as the article by Barro (1991). Rather it solves models with utility maximization given production constraints where human capital, learning by doing and comparative advantage in trade drive growth. He also alludes to, though never solves, an example where cities could cause growth by capturing certain agglomerations in production. This is the background philosophy for the model presented below: urbanization as a potential engine of growth. Although the model presented here does not solve the consumer maximization problem set out in Lucas, it does set out a production function which, if properly specified, could enter a growth model as a pertinent constraint. Further, it can be used to estimate individual cross-section urbanization agglomerations.

Section 2 of this paper presents the Cobb–Douglas production function model used in the paper. In this model urbanization is added as a shift factor. Section 3 presents the empirical results of the paper distinguishing between the dynamic results on the model including non-stationary variables, and the results from cross-section regressions. In particular, we apply the panel unit root test given by Im, Pesaran and Shin (1997); test of the null of cointegration in panel data from McCoskey and Kao (1998); and procedures for estimating long-run relationships from Pesaran and Smith (1995). Section 4 offers concluding notes.

II. THE MODEL

The model proposed here uses a Cobb–Douglas production function[1] which restricts the sum of exponents on capital and labor to one. The production function is defined for each country and each year:

$$y_{i,t} = A_{i,t}(U_{i,t})^{\lambda_i}(K_{i,t}^{\beta_i} N_{i,t}^{1-\beta_i}), \tag{1}$$

where $y_{i,t}$ is GDP for country i in time period t, $U_{i,t}$ is the percent of the population living in an urban area, $K_{i,t}$ is capital stock, and $N_{i,t}$ is the number of workers. $A_{i,t}$ is the specification for technology and is the element which introduces a stochastic element into the model. Technology and urbanization levels in this model both act as shift factors. Technology includes both a possible intercept and trend term:

$$A_{i,t} = e^{\alpha_i + \delta_i t + \varepsilon_{i,t}}. \tag{2}$$

[1]The Cobb–Douglas specification is the standard production specification used to estimate urban agglomerations in the urban literature. For example, see Henderson (1988, p. 34).

Normalizing by $N_{i,t}$ and taking the natural log forms the following model:

$$\ln y^*_{i,t} = \alpha_i + \delta_i t + \beta_i \ln K^*_{i,t} + \lambda_i \ln U_{i,t} + \varepsilon_{i,t}, \tag{3}$$

where $y^*_{i,t} = (y_{i,t}/N_{i,t})$ and $K^*_{i,t} = (K_{i,t}/N_{i,t})$.

The panel model here allows each of the cross-sections to have a unique intercept and time trend. Technology growth has a constant trend, δ_i, and a random shock, $\varepsilon_{i,t}$, that acts on the drift. β_i can be interpreted as the elasticity of capital per worker with respect to production output per worker and λ_i the urbanization elasticity. Allowing varying slopes as well as intercepts allows the individual countries to have heterogeneous production functions and means that the regression coefficients from each cross-section are estimated independently.

The data used in this paper come from the Penn World Tables and the World Bank *Social Indicators of Development*. The Penn World Tables provide yearly observations on non-residential capital stock per worker, KAPW, and real GDP per worker, GDPW. Both are reported in 1985 international prices. The World Bank *Social Indicators of Development* provides data on percent of the population living in an urban area. The data are recorded in yearly observations from 1965–1989. There are two potential groups of the data: developed and developing countries. The set of developing countries, Group 1, has 30 country observations and includes countries from Africa, Central America, South America and Asia. The set of developed countries, Group 2, has 22 observations comprised of European, Asian, North American countries, and Australia and New Zealand. As outlined in the model, all variables are in log form. All estimation and testing is done in GAUSS 3.0 using the package COINT 2.0.

With 25 years of observations for each cross-section series, we introduce a substantial time dimension which allows us to exploit current results in the time series and dynamic panel literature which allow us to test for unit roots and cointegration. The potential presence of unit roots and cointegrating relationships in the data opens up to us a whole new wealth of theory and interpretation unavailable in the cross-section.

III. EMPIRICAL RESULTS

3.1. Dynamic Panel Data Results

3.1.1. Testing for Non-Stationarity and Cointegration in Panel Data

In this section we summarize the non-stationary panel data tests for unit roots and cointegration we will be using and offer some intuition behind the testing. The test of the null hypothesis of cointegration states that under the H_0 there exists a long-run relationship between the natural logs of output per worker, capital stock per worker, and urbanization levels. The model allows for varying intercepts, trends and varying slopes and thus a coin-

TABLE 1
Panel Data Test Results

IPS Test			
Without a time trend			
Group 1	Urban	2.8469	Fail to reject
	KPW	−2.3542	Reject at 0.05
	GDPW	0.064	Fail to reject
Group 2	Urban	−1.3251	Fail to reject
	KPW	1.0377	Fail to reject
	GDPW	−0.2272	Fail to reject
With a time trend			
Group 1	Urban	−0.7235	Fail to reject
	KPW	n.a.	
	GDPW	4.0053	Fail to reject
Group 2	Urban	−2.3033	Reject at 0.05
	KPW	1.1549	Fail to reject
	GDPW	−1.6182	Fail to reject
Panel LM Test			
Group 1			
(GDPW, KAPW, Urban)	Without trend	−5.4516	Fail to reject
	With trend	−7.7522	Fail to reject
Group 2			
(GDPW, KAPW, Urban)	Without trend	−4.7427	Fail to reject
	With trend	−6.6709	Fail to reject
(KAPW, Urban)		8.4362	Reject at 0.01

tegration test for heterogeneous cross-sections is applicable. An intuitive interpretation of the null hypothesis would be that if there exists a long-run relationship between these three variables then including urbanization levels in the production function specification is reasonable and helpful in describing growth in output in the long run.

The first step in determining a potentially cointegrated relationship is to test whether the variables involved are stationary or non-stationary, i.e. whether the individual series contain unit roots. If all the variables are stationary, then traditional estimation methods can be used to estimate the relationship between the variables, in this case urbanization, KAPW and GDPW. If, however, at least one of the series is determined to be non-stationary then more care is required.

The test we use to test for stationarity was first presented by Im *et al.* (1997). In their paper, Im, Pesaran and Shin (IPS) present a statistic testing the H_0 of non-stationarity for a variable observed in a panel. The statistic is based on the augmented Dickey–Fuller (ADF) test widely used in the time series literature. Recall that the ADF test in the time series case can be written:

$$\Delta y_t = \alpha + \delta t + \rho y_{t-1} + \sum_{j=1}^{p} \gamma_j \Delta y_{t-j} + v_t, \tag{4}$$

where p denotes the number of lags. The null hypothesis of non-stationarity is written: $H_0: \rho = 0$ versus the $H_a: \rho < 0$. An equivalent way to express the null and alternative hypothesis is to use the notation: $H_0: y_i$ is an $I(1)$ process versus the $H_a: y_t$ is an $I(0)$ process. This null hypothesis can be tested using a type of t-statistic on ρ. However, because under the null hypothesis, this test statistic does not converge in distribution to a normal random variable, special tables of critical values have been constructed through extensive Monte Carlo simulation.[2]

In the panel case, the question is how to combine information on stationarity or non-stationarity for each individual cross-section into a conclusion about the panel as a whole. Assuming that the cross-sections are independent, IPS propose that the best way to combine information is to average the individual ADF t-statistics and use the following properties on the mean:

$$\Psi_{\bar{t}} = \frac{\sqrt{N}(\bar{t}_{N,T} - E[\bar{t}_{N,T}(p, 0)])}{\sqrt{\mathrm{Var}(\bar{t}_{N,T})}} \Rightarrow N(0, 1),$$

where $\Rightarrow$ denotes convergence in distribution, $\bar{t}_{N,T} = (1/N)\sum_{i=1}^{N} t_i$, t_i is the t-statistic for the *OLS* estimate of ρ in (4) for the ith unit of the cross-section, and $E[\bar{t}_{N,T}(p, 0)]$ is taken under the null hypothesis $\rho_i = 0$ for all i and with the choice $p = (p_1, p_2, \ldots, p_i, \ldots, p_N)'$ of the lag-length vector for the regressions unit by unit in (4). $\Psi_{\bar{t}}$ can be compared to critical values for a one-sided $N(0, 1)$ distribution. The moments of $\bar{t}_{N,T}$ depend on the number of time series observations and the appropriate lag order, p_i, for each cross-section. IPS provide the necessary tables to construct these moments for each individual data set. The selection of the appropriate lag order for the variables here follows the procedure suggested by Campbell and Perron (1991).

If we find that GDPW and one or both of the variables KAPW and urbanization are non-stationary, then we can test the system for cointegration. The residual-based test for cointegration we use comes from McCoskey and Kao (1998). The test is constructed from the partial sums of the estimated residuals of a regression equation of non-stationary variables. It is a panel data version of the LM-statistic proposed by Harris and Inder (1994). The precise form of the test is given:

[2]For a summary of tests and appropriate critical values, see Hamilton (1994, pp. 528–529, 762–763) or Fuller (1976, pp. 371–373).

$$\overline{LM} = \frac{1}{N}\sum_{i=1}^{N}\left(\frac{(1/T^2)\sum_{t=1}^{T} S_{i,t}^{+2}}{\hat{\varpi}_{i1.2}^2}\right),$$

where $S_{i,t}$ is the partial sum of estimated residuals,

$$S_{i,t}^{+} = \sum_{j=1}^{t} \hat{\varepsilon}_{ij},$$

and the residuals, $\hat{\varepsilon}_{i,t}$, can be estimated either using fully modified (*FM*) estimation or dynamic *OLS* (*DOLS*) in equation (3).[3] $\hat{\varpi}_{i1.2}^2$ is a consistent estimator of $\varpi_{i1.2}^2$ where $\varpi_{i1.2}^2$ is defined in McCoskey and Kao (1998) and $\hat{\varpi}_{i1.2}^2$ can be found easily in econometric packages, e.g., COINT 2.0.

It can be shown that, for example, McCoskey and Kao (1998):

$$LM^* = \sqrt{N}(\overline{LM} - \mu_v) \Rightarrow N(0, \sigma_v^2), \tag{5}$$

where $\mu_v = E[\int V^2]$, $\sigma_v^2 = \text{Var}[\int V^2]$ and $\int V^2$ is defined in McCoskey and Kao. Thus, an appropriately normalized version of the statistic converges to a normally distributed random variable with mean zero. In this limiting distribution, μ_v and σ_v^2 are the mean and variance, respectively, of a complex functional of Brownian motion. These values depend only on the number of regressors. The appropriate values for up to five regressors can be found in McCoskey and Kao. Although McCoskey and Kao allow only for the specification where $\delta_i = 0$, we extend their results to allow for a possible drift term in the cointegrating relationship.[4] The derivation of this more general case in the strict time series dimension can be found in Shin (1994).

Essentially, this test is combining evidence from averaging the *LM*-statistic across the cross-sections. The test is one-sided: large values of LM^* correspond to estimating non-stationary residuals and will result in rejection of the null hypothesis of cointegration (equivalent to rejecting the stationarity of the errors). Rejection of LM^* concludes that the average of the individual *LM*-statistics across the countries in the panel is far away from the mean, μ_v, constructed under the null.

3.1.2. Unit Root Test Results

In this section, we test the series for stationarity. In the first case we assume that none of the individual series in our model contains a trend. Thus, it is assumed for each series, $y_{i,t}$ that $E(\Delta y_{i,t}^*) = 0$. This means that each series

[3]There is some evidence that the *FM* method may be more powerful in small data sets but the demands on the data are also greater. Where possible the *FM* estimation is used and is noted specifically. However, where the tested cointegrating relationship contains two regressors, the *DOLS* method is used including two leads and two lags in the estimation.

[4]For the extension we found the values 0.04714 and 0.00083 for μ_v and σ_v^2 through simulation. These values are for the case with two regressors.

could contain a non-zero intercept but not a time trend. To test the three series of urbanization, KAPW and GDPW for stationarity in our panels of developed and developing countries, we can use the *ADF* test given in equation (4) where $\delta_i = 0$ to construct the appropriate $\Psi_{\bar{t}}$.

The $\Psi_{\bar{t}}$ for this set of pooled results are as follows:

	Urbanization	*KAPW*	*GDPW*
Group 1	2.8468830	−2.3542323	0.063962916
Group 2	−1.3251332	1.0377193	−0.22716037

As it is a one-sided test, a statistic less than −1.645 would cause rejection of the null of non-stationarity. The only series which would reject the null is capital stock per worker in developing countries.

This result is quite interesting and underlines the necessity of dividing the countries into two groups. While both groups have GDPW and urbanization levels which seem to be continually growing over time, only the set of developed countries experiences such growth in capital per worker. This result seems to agree nicely with what we might predict about one of the fundamental differences between developing and developed countries: the ability to accumulate capital.

However, our assumption that $\delta_i = 0$ may be overly restrictive, especially in the case of GDPW; for example, see Canjels and Watson (1997). Therefore we test stationarity again allowing for a time trend. (As KAPW for Group 1 has already rejected the null of non-stationarity without a time trend, we do not test it here.)

	Urbanization	*KAPW*	*GDPW*
Group 1	−0.7235	*X*	4.0053
Group 2	−2.3022	1.1549	−1.6182

In this case, only urbanization for Group 2 can reject trend stationarity at the 5 percent level. This adds another interesting economic interpretation to the results: urbanization in developed countries is occurring at a stable, steady rate although the same cannot be said for the set of developing countries. Again this seems to agree with observed differences between the two groups. Given the presence of non-stationary variables in both specifications, we now proceed to test for cointegration.

3.1.3. Results for Cointegration (Without Trend)

The first step in investigating a cointegrating relationship is to be sure that the regressors themselves are not cointegrated. The theory of testing for cointegration is applicable only under the assumption that the independent variables themselves are not cointegrated. Therefore, the next step is to test for a cointegrating relationship between the variables $\ln K^*_{i,t}$ and $\ln U_{i,t}$. For the regression with Group 1, the variables cannot be cointegrated because the natural log of capital per worker is stationary. However, with Group 2 the two variables may be cointegrated and a residual-based test must be done. The following relationship is tested:

$$\ln U_{i,t} = \psi_i + \delta_i \ln K^*_{i,t} + u_{i,t}.$$

This test is constructed using μ_v and σ_v^2 equal to 0.1162 and 0.0109, respectively, from McCoskey and Kao (1998) for construction of the test with one regressor. The null hypothesis of cointegration between capital per worker and urbanization is rejected for Group 2 with $\overline{LM} = 8.4362$, far away from the one-sided critical value of 1.645.[5]

Now we can test the cointegrating relationship:

$$\ln y^*_{i,t} = \alpha_i + \beta_i \ln K^*_{i,t} + \lambda_i \ln U_{i,t} + \varepsilon_{i,t}.$$

Because the test for cointegration is based on estimated residuals, the parameter estimates are reported simultaneous to construct the test statistic. The individual *LM* test statistics are reported in Tables 2 and 3 and the individual parameter estimates are reported in Tables 6 and 7, below.

Originally the test of the null hypothesis of cointegration, *LM*, was proposed for use in the literature as the test of the null of no cointegration was thought to have low power. Therefore, it may be useful to check the individual time series results against the individual *ADF* test statistics for the null of no cointegration. The *ADF* test of the null of no cointegration is analogous to the *ADF* test for unit roots from equation (4), except that the test is now based on estimated residuals from the estimated equation. Note the *ADF* test for cointegration is constructed without intercept and trend. The critical values of the test depend on the estimation and are no longer the same as those for the unit root test. These individual *ADF* results are also reported in Tables 2 and 3. For no country was the null hypothesis of cointegration rejected, although the null hypothesis of no cointegration was rejected six times for Group 1 and once for Group 2.

It is clear from the results that low power is an issue with both the time series results for the *LM* and *ADF* tests for cointegration. In fact in only a few cases can the null be rejected with either test. The low power of the tests is a major motivation for pooling data into a panel. When we pool our results, for Group 1, the null hypothesis of cointegration cannot be rejected with $LM^* = -5.416$[6].

Consider the parameter estimates for Group 1. As the LM^* test statistic has failed to reject the null hypothesis of cointegration, the vector of estimated coefficients can be interpreted as the potential cointegrating vector of the system. These estimates can be interpreted as long-run impacts. Considering the estimated parameters in Table 6, it is encouraging

[5]In this case, with one regressor, estimation was done with the fully modified procedure.

[6]μ_v and σ_v^2 equal to 0.0850 and 0.0055, respectively, are used as the mean and variance for two regressors and are reported in McCoskey and Kao (1998).

TABLE 2
Individual LM and ADF Cointegration Test Results: Group 1

Without trend	*LM test*	*ADF test*
Argentina	0.0165	−3.7610*
Bolivia	0.0060	−3.1589
Chile	0.0160	−3.2163
Colombia	0.0164	−2.4666
Côte D'Ivoire	0.0079	−2.2913
Dominican Republic	0.0190	−1.7625
Ecuador	0.0153	−3.1520
Guatemala	0.0082	−2.2419
Honduras	0.0074	−2.0679
India	0.0075	−1.2656
Iran	0.0102	−2.2754
Jamaica	0.0098	−1.5712
Kenya	0.0084	−4.0187**
Madagascar	0.0121	−3.4718*
Malawi	0.0073	−3.6066*
Mexico	0.0204	−3.0512
Morocco	0.0087	−2.6700
Nigeria	0.0130	−1.7814
Panama	0.0085	−2.2628
Paraguay	0.0053	−3.5590*
Peru	0.0158	−1.4895
Philippines	0.0295	−1.9184
Sierra Leone	0.0058	−3.3807
Sri Lanka	0.0098	−5.3311**
Syria	0.0087	−0.5201
Thailand	0.0096	−2.7996
Turkey	0.0079	−1.8790
Venezuela	0.0069	−0.8383
Zambia	0.0093	−3.0685
Zimbabwe	0.0084	−2.6693

Notes:
(a) Lag orders calculated for each country according to Campbell and Perron (1991).
(b) Critical values at the 5 (**) and 10 (*) percent level for the *LM* test: 0.2177 and 0.1617 from Harris and Inder (1994).
(c) Critical values at the 5 (**) and 10 (*) percent level for the *ADF* test: −3.7675 and −3.4494 from Phillips and Ouliaris (1990).

to note that, for most cases, the coefficient β_i, which can be interpreted as the elasticity of output per worker with respect to capital per worker, is positive. In almost two-thirds of the cases, the estimated parameter lies between 0 and 1 which is consistent with the model specification, although 10 of the 16 significant estimates are greater than 1. Only Peru has an estimate of the elasticity of output per worker with respect to capital per

TABLE 3

Individual LM and ADF Cointegration Test Results: Group 2

Without trend	*LM test*	*ADF test*
Australia	0.0096	−1.9237
Austria	0.0105	−2.1578
Belgium	0.0055	−1.5610
Canada	0.0158	−2.4516
Denmark	0.0083	−2.4228
Finland	0.0098	−2.8373
France	0.0127	−2.5032
Federal Republic Germany	0.0188	−3.1131
Greece	0.0156	−2.7878
Ireland	0.0050	−3.0782
Italy	0.0124	−3.3764
Japan	0.0059	−2.8322
Korea	0.0070	−1.9803
Luxembourg	0.0102	−2.7900
Netherlands	0.0099	−3.4491
New Zealand	0.0108	−2.7137
Norway	0.0079	−1.9528
Spain	0.0079	−1.5837
Sweden	0.0110	−1.5840
Switzerland	0.0052	−3.0540
United Kingdom	0.0126	−2.9731
USA	0.0077	−3.8969**

Notes:

(a) Lag orders calculated for each country according to Campbell and Perron (1991).

(b) Critical values at the 5 (**) and 10 (*) percent level for the *LM* test: 0.2177 and 0.1617 from Harris and Inder (1994).

(c) Critical values at the 5 (**) and 10 (*) percent level for the *ADF* test: −3.7675 and −3.4494 from Phillips and Ouliaris (1990).

worker which is negative and significant. It is also interesting to note that exactly half, or 15 of the 30 countries, have negative estimates for λ_i. Fifteen of the 30 estimates for λ_i are significant with seven of the 15 less than 0 and eight greater than 0. A negative λ_i would mean a negative elasticity of output with respect to urbanization. The result is interesting within the context of the debate of whether or not developing countries have become over-urbanized.[7]

[7]If we reject cointegration then we encounter the problem of estimating a spurious regression. As discussed in Granger and Newbold (1974) and Phillips (1986), a spurious regression of two independent non-stationary series will tend to show a significant relationship when none exists. The problem gets worse as the time dimension increases. In the absence of a cointegrating relationship, the specification is spurious. A spurious regression has the following characteristics: (a) estimates are not consistent and converge to random variables, not constants; (b) *OLS* t and F statistics diverge; (c) R^2 may not tend to 0. Thus, caution is suggested when interpreting results from spuriously estimated regressions.

The results for the testing on Group 2 are similar: the null hypothesis of cointegration cannot be rejected with $LM^* = -4.7427$. Results for individual testing, the *LM* and *ADF* test statistics are provided in Table 7 for Group 2. Parameter estimates are provided in Table 7. In this group, 10 out of the 22 estimates for urbanization are significant with only three of those less than 0. For the estimates on KAPW, 13 out of 22 are significant with none of the significant estimates less than 0 and only four greater than 1.

Given our production function specification, there is an important link between λ_i, β_i and output per worker. In particular, for countries where $\lambda_i < 0$, there is an added imperative for capital accumulation in order to see growth in output per worker. Starting with the production function

$$y^*_{i,t} = A_{i,t}(U_{i,t})^{\lambda_i}(K^{*\beta_i}_{i,t})$$

and holding $A_{i,t}$ constant, we can take the total differential with respect to U and K^*:

$$dy^*_{i,t} = A_{i,t}\lambda_i(U_{i,t})^{\lambda_i - 1}(K^{*\beta_i}_{i,t})dU + A_{i,t}(U_{i,t})^{\lambda_i}\beta_i(K^{*\beta_i - 1}_{i,t})dK^*.$$

We set $dy^*_{i,t} = 0$ to find the trade-off between dU and dK^* necessary to hold output constant. It can be shown that

$$dK^* = -\left(\frac{\lambda_i}{\beta_i}\frac{K^*}{U}\right)dU.$$

It can be seen directly, and is quite intuitive, that larger negative values of λ_i require larger levels of change in capital accumulation per worker for economic growth. Thus, for developing countries, over-urbanization has a very similar effect as high birth rates in terms of future economic growth. For countries where $\lambda_i > 0$, countries can experience decreases in capital per worker and yet still experience growth in output through the positive effects of urbanization.

How should one interpret the failure to reject the null hypothesis of cointegration? Statistically speaking, the failure to reject the null means that we cannot rule out a long-run relationship between the natural log of GDPW, KAPW and urbanization levels for either developed or developing countries. However, given the potential importance of a time trend to the underlying specification, in the next section we test for cointegration allowing for a non-zero time trend.

3.1.4. Results for Cointegration (With Trend)

In the previous section we assumed that none of the series contained a time trend and that the regression itself contained no trend. In this section, we allow for the presence of a time trend and test for cointegration of the relationship:

$$\ln y^*_{i,t} = \alpha_i + \delta_i t + \beta_i \ln K^*_{i,t} + \lambda_i \ln U_{i,t} + \varepsilon_{i,t}.$$

For this case we could not reject cointegration for either group. Group 1 had $LM^* = -7.7522$ and for Group 2, $LM^* = -6.6709$. The individual country results for the LM and ADF tests are given in Tables 4 and 5. The parameter estimates are given in Tables 6 and 7.

The first feature of this model to investigate is the significance of the

TABLE 4
Individual LM and ADF Cointegration Test Results: Group 1

With trend	*LM test*	*ADF test*
Argentina	0.0092	−4.3228**
Bolivia	0.0047	−3.9331*
Chile	0.0113	−3.7545
Colombia	0.0151	−2.5185
Côte D'Ivoire	0.0064	−2.5648
Dominican Republic	0.0051	−2.5088
Ecuador	0.0048	−3.4957
Guatemala	0.0064	−2.3857
Honduras	0.0038	−2.1421
India	0.0034	−2.1467
Iran	0.0054	−2.8751
Jamaica	0.0050	−1.5862
Kenya	0.0070	−3.8091
Madagascar	0.0164	−3.4908
Malawi	0.0060	−3.6405
Mexico	0.0051	−2.8755
Morocco	0.0046	−2.9733
Nigeria	0.0051	−2.9910
Panama	0.0034	−2.2657
Paraguay	0.0035	−3.9425
Peru	0.0046	−2.8547
Philippines	0.0044	−2.8913
Sierra Leone	0.0041	−4.1300*
Sri Lanka	0.0047	−4.3420**
Syria	0.0073	−1.3797
Thailand	0.0069	−3.1136
Turkey	0.0074	−3.4411
Venezuela	0.0053	−2.0456
Zambia	0.0065	−3.1550
Zimbabwe	0.0066	−2.6623

Notes:
(a) Lag orders calculated for each country according to Campbell and Perron (1991).
(b) Critical values at the 5 (**) and 10 (*) percent level for the *LM* test: 0.101 and 0.081 from Shin (1994).
(c) Critical values at the 5 (**) and 10 (*) percent level for the *ADF* test: −4.1567 and −3.8429 from Phillips and Ouliaris (1990).

TABLE 5
Individual LM and ADF Cointegration Test Results: Group 2

With trend	*LM test*	*ADF test*
Australia	0.0082	−2.2845
Austria	0.0073	−2.8297
Belgium	0.0039	−1.6244
Canada	0.0041	−2.0781
Denmark	0.0058	−2.9091
Finland	0.0035	−2.9431
France	0.0094	−3.2818
Federal Republic Germany	0.0045	−3.2896
Greece	0.0092	−3.6493
Ireland	0.0041	−2.9318
Italy	0.0054	−3.3389
Japan	0.0053	−3.0194
Korea	0.0064	−3.4162
Luxembourg	0.0061	−3.0240
Netherlands	0.0045	−2.8876
New Zealand	0.0032	−2.9273
Norway	0.0069	−2.3047
Spain	0.0053	−1.5910
Sweden	0.0084	−1.5066
Switzerland	0.0048	−3.0625
United Kingdom	0.0109	−3.3169
USA	0.0070	−3.8055

Notes:
(a) Lag orders calculated for each country according to Campbell and Perron (1991).
(b) Critical values at the 5 (**) and 10 (*) percent level for the *LM* test: 0.101 and 0.081 from Shin (1994).
(c) Critical values at the 5 (**) and 10 (*) percent level for the *ADF* test: −4.1567 and −3.8429 from Phillips and Ouliaris (1990).

estimated trend. The trend has an important economic meaning; it can be seen as the constant growth rate in GDPW caused by factors other than urbanization and capital per worker. Again, there is a noticeable difference between the two groups. In Group 1, in 11 out of the 30 cases, the trend is significant. In six out of these 11 cases, the estimate is negative. Of these significant estimates, the maximum (Jamaica) is 0.9061 and the minimum (Morocco) −1.9375. In the case of Morocco, this negative trend is offset by a significant and dramatic impact of urbanization. With Group 2 countries, eight countries have a significant time trend and, of these, six are positive. Only Greece and the Netherlands have negative and significant growth rates. The maximum, significant, growth rate estimate is 0.1747 (Luxembourg) and the minimum is −0.651 (Greece).

TABLE 6

Parameter Estimates (Potential Cointegrating Vector): Group 1

	Without trend		*With trend*		
	Urbanization	*KPW*	*Trend*	*Urbanization*	*KPW*
Argentina	−11.6618	0.8650	−0.1064	−0.7074	1.1800
Bolivia	1.1361	0.5123**	0.0218	−0.6778	0.3325
Chile	−1.2318	2.4682**	−0.3616	30.4351	−1.9892
Colombia	11.7858	0.1485	−0.3479**	29.5188*	−1.5471
Côte D'Ivoire	−0.0525	0.7477	0.0466	−1.8189	0.9726
Dominican Republic	−5.5952	0.9790	−0.2453**	7.3350	0.3406
Ecuador	−4.9294**	2.4941**	−0.0699**	−2.8854	3.4337**
Guatemala	1.2456**	0.6837**	0.0058	1.0373**	0.5986**
Honduras	4.5996**	0.6486**	−0.1956	14.5110**	0.2431
India	2.2072**	−0.4081	0.7104**	−32.2889**	−3.5229**
Iran	6.7274	−0.6880	−0.7079**	63.2460**	−6.7503**
Jamaica	1.2617	1.1434**	0.9061**	−74.7119	2.5179**
Kenya	0.4166*	0.0274	−0.0680	1.9980	0.01552
Madagascar	−0.2229*	0.2239*	0.5450	−19.6198	0.0890

Malawi	0.6876**	0.1084	−0.2274	5.9497	−0.1473
Mexico	−7.0563**	0.6170**	0.2377**	−30.0058**	2.2319**
Morocco	2.3444**	−0.4562	−1.9375**	117.5029**	−1.9094**
Nigeria	−1.4528	2.6991**	−1.9047	63.2119	0.7985
Panama	−0.9437	1.6904**	0.1348**	−15.0350**	0.6024
Paraguay	−1.9483*	1.0581**	0.1181	−10.2285	1.0989**
Peru	−0.1646	−2.2677**	0.1755**	−22.7104**	−4.3067**
Philippines	0.8339	0.7503	−0.3930**	36.5182**	−0.4431
Sierra Leone	−3.5124**	1.5436**	−0.0209	−1.7616	1.3988**
Sri Lanka	−9.4687**	0.4426	−0.0242	−11.3452**	1.2327**
Syria	−0.7261	1.0406	−0.8786	94.6113	2.2382
Thailand	0.6625	0.3681	0.4593	−15.8568	−0.0570
Turkey	1.9848**	0.4513	0.0217	0.7135	0.2024
Venezuela	−5.4436**	2.3985**	0.0118	−6.6991**	2.5117**
Zambia	−0.4664	0.4665	−0.1652	3.6718	−0.2418
Zimbabwe	4.7156**	2.7276**	0.0967	0.4574	2.6352**

Notes:
(a) Significant at 10 (*) and 5 (**) percent was determined using critical values (in absolute value) 1.645 and 1.96.
(b) Model estimated in COINT 2.0.

TABLE 7
Parameter Estimates (Potential Cointegrating Vector): Group 2

	Without trend		With trend		
	Urbanization	KPW	Trend	Urbanization	KPW
Australia	6.4541	0.8745**	−0.4452	1.3855	1.9614
Austria	−2.4958	0.7775**	1.5074	−43.5951	1.2120**
Belgium	−1.570	0.8314**	−0.9334	13.0072	1.0748**
Canada	19.1272**	0.1248	0.1040**	84.2047**	−4.6807**
Denmark	5.3582**	−0.0481	0.1168**	−48.9148*	4.0053**
Finland	0.6886**	0.5484**	−0.0312	−0.8065	1.7441**
France	1.3130	0.3278	−0.0140	0.7786	0.9982
FRG	−13.2439**	0.5286**	0.0978**	−68.3376**	1.7803**
Greece	0.5740	0.3298	−0.6591*	81.3123*	0.6687*
Ireland	5.1814**	−0.1106	−0.1054	15.8632	0.8442
Italy	−7.8037**	1.9489**	0.0440	−13.7825**	0.9990**
Japan	−0.4680	0.6040**	−0.0074	−1.4348	0.7712
Korea	0.1000	0.9230**	0.0233	−1.0339	1.0663
Luxembourg	1.0607	1.2684**	0.1747**	−5.6439**	−2.3241
Netherlands	9.2620**	−0.0537	−0.0215**	12.2456**	0.7009*
New Zealand	8.2611**	0.0065	0.0207**	11.3555	−1.4035**
Norway	−0.3109	2.0965**	−0.0212	1.5176	2.9486
Spain	−3.2181**	1.0908**	0.0877*	−16.6840**	1.1093**
Sweden	8.0661**	0.0985	0.0060	7.0590	−0.1800
Switzerland	−0.5293	0.4914**	0.0533	−3.9242	−0.2440
United Kingdom	2.0313	0.4612**	−0.0516	37.3485	−0.7082
USA	−9.6019	0.2910	0.0284	−45.1144	−0.5258

Notes:
(a) Significant at 10 (*) and 5 (**) percent was determined using critical values (in absolute value) 1.645 and 1.96.
(b) Model estimated with intercept in COINT 2.0.

There is an interesting interaction between urbanization and the time trend. In 15 cases across both groups, both urbanization and the time trend are significant. In all but one country (Canada) the signs are reversed. It seems that urbanization is either draining from otherwise positive growth or propping up growth rates in an otherwise struggling country. In either case, urbanization seems crucial in understanding future output. For Group 1, four countries have positive urban elasticities combined with negative growth rates (Colombia, Iran, Morocco, and the Philippines); in the remaining five cases a negative urban elasticity is combined with a positive growth rate (India, Jamaica, Mexico, Panama, and Peru). For Group 2, two countries have a positive urban elasticity combined with negative growth rate (Greece and the Netherlands) and four countries have a positive growth rate combined with negative urban elasticity (Denmark, FRG, Luxembourg, and Spain.) Overall, including the trend in the specification results in much more dramatic estimates for urbanization.

In comparing results for urbanization across the two specifications, in almost all cases both signs remained the same. There were two glaring exceptions: the estimate of urban elasticity for Denmark changed from 5.3583 without a trend to -48.9148 with a trend; for India the estimate changed from 2.2072 without a trend to -32.2889. The results for capital per worker remained stable across the two models. In the model with trend, in Group 1 there are 13 significant estimates with four of these negative. In the case of Group 2, 11 estimates were significant with two of these less than zero.

3.2. *Cross-Section Results and Average Effects*

3.2.1. *Testing for Average Effects*

The results from the above cointegration analysis should be considered in contrast to more traditional cross-section results. In fact, the two types of studies are answering very different questions. The cross-section studies traditionally attempt to find average long-run effects rather than examining specific paths of different countries. In Pesaran and Smith (1995) two methods are given to consistently estimate these long-run averages in the presence of cointegrating relationships.

The first method is given as simply taking the average across the individual parameter estimates for each cross-section. Thus:

$$\overline{\beta} = \sum_{i=1}^{N} \beta_i.$$

The second suggested method involves estimating a cross-section relationship between the averages, across time, of the groups. Thus the estimated regression is defined as:

$$\overline{y}_i^* = \alpha + \beta \overline{K}_i^* + \lambda \overline{U}_i + \varepsilon_i,$$

where $\overline{y}_i^* = \sum_{t=1}^{T} \ln y_{i,t}^*$, $\overline{K}_i^* = \sum_{t=1}^{T} \ln K_{i,t}^*$, and $\overline{U}_i = \sum_{t=1}^{T} \ln U_{i,t}$.

In this specification, the estimates will be consistent. However, the usual standard errors are not valid. Pesaran and Smith suggest White's heteroscedasticity-consistent standard errors. It should be noted that using the cross-section approach to estimate average effects has one major disadvantage over the dynamic approach: in this approach the regressors are assumed to be strictly exogenous. Both the dynamic and cross-section approaches require independence across the cross-sections for the asymptotic results to hold.

Pesaran and Smith do provide an important caution for those studies where very short time periods are used to estimate average effects (such as the strict cross-section approach), such estimations are likely to be biased or inconsistent. Thus, even in this 'cross-section' approach, the time dimension of the panel is crucial.

3.2.2. *Results for Average Effects*

Average of the Estimated Parameters Following the guidelines by Pesaran and Smith, our first method of averaging the individual estimated coefficients for urbanization and KAPW across the cross-sections yields the following results:

	Group 1	*Group 2*
Urbanization	−0.4760	1.2835
	(1.2856)	(1.0291)
KAPW	0.7828	0.6096
	(0.1442)	(0.1181)

The standard errors reported are calculated under the assumption that the cross-sections are independent and using the usual properties on variances of the average of independent random variables.

At first glance it is disappointing to realize that for neither group is the coefficient on urbanization further than two standard deviations away from 0; however, when considering the estimates on which this is based such an inconclusive result is predictable. When the original estimates were done in both groups the sample was almost split with both positive and negative coefficients on urbanization. The result on KAPW is much clearer—in both cases the coefficient is positive and at least two standard deviations away from zero.

Average Cross-Section Results Using the second method suggested and constructing averages across time and using *OLS*, we obtain the following results:

	Group 1	*Group 2*
Urbanization	0.8954	0.5519
	(0.1464)	(0.2020)
	(0.9933)	(7.2275)
KAPW	0.1668	0.5479
	(0.0597)	(0.0860)
	(0.2717)	(2.9682)

The first set of standard errors reported are those from the original *OLS* estimation; the second set of standard errors reported are calculated from White's heteroscedasticity-consistent estimator variance−covariance matrix.

The results from this estimation, when taken with the second set of standard errors, coincides strongly with the previous results with regards to urbanization. In neither case is urbanization significantly different from 0. However, in this estimation, the results on KAPW are also inclusive.

It is clear that using White's heteroscedasticity-consistent estimator corrections for the standard errors makes an enormous difference in the interpretation of the results, especially in the case of Group 2. The *OLS* standard errors would cause us to conclude that all estimates are positive and significantly different from 0. It is interesting to note that, despite the inconclusive results on the parameters, the estimated R^2 for Group 1 is 0.848 and for Group 2 it is 0.787.

When taken as a whole, these results seem to support that average effects of the elasticity of capital per worker with respect to output to worker are positive. This result is nice but not very groundbreaking. With regard to urbanization, the results are much less conclusive. In fact based on the inability of the estimates to determine even the sign on the average effect of the elasticity of urbanization we conclude that attempting to estimate such an average effect may be misguided—based on our dynamic studies it seems clear that the impact of urbanization varies greatly across the cross-sections. Such an individualized impact, which may depend crucially on internal economic and political structure, can be best captured in the dynamic, heterogeneous approach.

IV. CONCLUSION

Urban economists have been anxious to link output and measures of urbanization. Cities form, it is assumed, in response to market forces in production. In this essay an attempt is made to pin down the exact relationship between urbanization and output over time using results from the time series literature and non-stationary panel data literature. A traditional constant returns to scale Cobb–Douglas production function is specified with urbanization as a shift factor. The results show clearly that this specification cannot be rejected and may be useful in understanding long-run growth. However, even if urbanization is crucial to growth, our results show that the impact of urbanization varies greatly across countries and therefore attempts to identify or even determine the sign of constant long-run effects are misguided. Understanding whether urbanization is an engine or anchor to growth is crucial for policy in both developing and developed countries.

There are other important results from this study as well. Testing presented here shows that the natural logs of urbanization and GDP per worker are non-stationary for the group of developing countries and for developed countries, the natural logs of GDP per worker, capital per worker and urbanization are all non-stationary. Serious studies of the dynamic relationship between growth in GDP and urbanization should take heed of these results; otherwise, estimation may be spurious. In addition, using results from Pesaran and Smith, it should also be mentioned that estimating average long-run effects with a simple cross-section may be biased or inconsistent. It is clear that using results from the time series literature and dynamic panel literature may greatly facilitate our understanding of the intersection of urbanization and growth.

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