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INTER-SECTORAL TERMS OF TRADE AND INVESTIBLE SURPLUS

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Abstract

This paper studies the validity of the Preobrazhensky's First Proposition, P1, for the centrally planned Poland during the period of 1960-1987 by testing whether the state increased its internal accumulation, investible surplus, by reducing the inter-sectoral terms of trade between agriculture and industry. It uses an alternative bivariate approach with pair-wise cointegration and Granger causality analysis. While finding some support for P1 in Poland, it reveals that the existing multivariate empirical approach derived from the Sah-Stiglitz's market-based theoretical model of the 'price scissors' problem can misrepresent not only the validity of P1 in a centrally planned economy but also the impact of the intrinsic determinants of the state's investible surplus such as the production capacities in agriculture and industry.

JEL Code: O2, P2

Keywords: Price Scissors, Internal Accumulation, Investible Surplus, Inter-sectoral Terms of Trade, Central Planning, Poland

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1. Introduction

Since the famous industrialization debate in the Soviet Union during the 1920s, a persistent question in development economics has been whether reducing the price of agricultural to industrial goods, the inter-sectoral terms of trade, can help extract a surplus from agriculture for industrialization in the early stages of economic development.[1] This issue has been crucial not only for the centrally planned economies but also for the developing countries dependent on agriculture. The most prominent advocate of this strategy for the Soviet Union was Evgeny Preobrazhensky (1926, translated 1965) who argued that the state could increase its internal accumulation, investible surplus, by reducing the inter-sectoral terms of trade against agriculture, in favor of the industry. This argument is often referred to as the Preobrazhensky's First Proposition, P1.[2] However, the validity of this important proposition was not investigated theoretically or empirically till the 1980s.

Sah and Stiglitz (1984) provided a first theoretical proof of P1 in a general equilibrium dualistic model depicting a two-sector (rural-urban) economy, but under restrictive assumptions. Shortly after, the robustness of their proof was called into question by, among others, Li and Tsui (1985), Carter (1986), Blomqvist (1986),[3] and by Sah and Stiglitz themselves (1986, 1987a, 1987b).[4] This theoretical debate culminated in

[1] The study of this question has given rise to many studies and books resulting in a growing literature under the name of 'price scissors' economics.

[2] Preobrazhensky's second proposition covers the distributional consequence of changing the inter-sectoral terms of trade excluded in this study. It states that, by lowering the agriculture's terms of trade, it is possible to increase the investible surplus without lowering the economic position of the industrial workers. It, however, does not hold in the Sah and Stiglitz's original model.

[3] Li and Tsui (1985) showed that P1 might not hold in the Sah-Stiglitz framework if urban/industrial workers behaved according to the efficiency wage hypothesis. Carter (1986) raised the same point as well as the unsettling impact of open-economy considerations on P1. On the other hand, Blomqvist (1986) showed that Preobrazhensky's propositions hold under a rural-urban price differential for agricultural goods.

[4] It should be noted that because of its primary focus on the 'price scissors' mechanism the Sah-Stiglitz's

a book by Sah and Stiglitz (1992) that provided a comprehensive analytical framework for various aspects of the 'price scissors' problem,[5] making the validity of P1 for any particular country in effect an empirical matter.

This paper analyzes the validity of the Preobrazhensky's First Proposition, P1, for the centrally planned Poland during the period of 1960-1987 when the state used the inter-sectoral terms of trade for extracting a surplus from its mostly private-farm agriculture for industrialization.[6] It discusses the deficiencies of the existing multivariate empirical approach derived from the Sah-Stiglitz's market-based theoretical framework and of the commonly used traditional estimation methods for testing the validity of P1 in the centrally planned economies. It proposes and applies to the Polish data an alternative bivariate empirical approach and time-series estimation methods a' la pair-wise cointegration and Granger causality. It reveals that the existing multivariate approach combined with the traditional estimation methods can misrepresent not only the validity of P1 but also the impact of the intrinsic determinants of the state's investible surplus such as the production capacities in agriculture and industry.

The rest of the paper is organized as follows. Section 2 presents the underlying theoretical framework, empirical methodology, and data sources of this study. Section 3 tests the validity of the Preobrazhensky's First Proposition, P1, for Poland during the

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models could not provide a theoretical underpinning for studying the comprehensive patterns of the 'inter-sectoral resource flows' in the development process. For details and a survey of this broader literature, see Karshenas (1995); and for a survey of the empirical studies on the direction and size of the net resource flow between agriculture and industry in China, see Sun (2000).

[5] This was not the end of criticism of the Sah-Stiglitz's model, or the theoretical innovations and extensions of the framework used for analyzing the 'price scissors' problem. For example, Baland (1993) added demand rationings to the model and considered the possibility of agricultural exports in the case of an excess supply of the rural good. Knight (1995) used offer curve analysis from trade theory to clarify and illuminate aspects of the previous work and then applied his new approach to the case of China illustratively. Sun (2000) extended the framework by incorporating the production and trade of (some) industrial consumer goods within the rural sector and considered demand rationing.

[6] Private farms constituted 70% of the land under cultivation. For an overview of the Polish economy during this period, see Advocate (2008).

period 1960-87 using an alternative bivariate approach along with stationarity, cointegration, and Granger causality analyses. Section 4 presents some concluding remarks.

2. Theoretical Framework, Empirical Methodology, and Data Sources

Existing empirical studies of the 'price scissors' problem have used variations of the original Sah-Stiglitz's model to motivate a multivariate relationship between the investible surplus per capita, *IS*, as the dependent variable, and its determinants as the independent variables. For a closed-economy case, the terms of trade between agriculture and industry, *TT*, and the lagged capital-labor ratios as a measure of the 'production capacity' in these two sectors, *KA(-1)* and *KI(-1)*, are used as the independent variables (Li-Tsui (1985, 1990)). For an open-economy case, the net exports per capita,[7] *NE*, is added as a fourth independent variable (Advocate (2008)). The inclusion of *KA(-1)*, *KI(-1)*, and *NE* in the specification is supposed to control for their impact on *IS* and, hence, to isolate the impact of *TT* on *IS*.[8] The terms of trade, *TT*, is set by the state and is an exogenous policy tool.[9]

The theoretically expected impacts on *IS* of its main determinants are as follows. The impact of *TT* would be negative or positive depending on whether the much-debated Preobrazhensky's First Proposition, P1, holds or not respectively. The impact of *KA(-1)* could be positive or negative. It would be positive if the resulting increase in the supply

[7] This excludes any net revenue (subsidy) that the government may derive (pay) because of the difference between the domestic and international terms of trade, related to the net imports of food.
[8] While Li-Tsui (1985, 1990) finds no support for P1 in China for the period 1952-82, Advocate (2008) finds strong support for P1 in Poland for the period 1960-87.
[9] While studies of the 'price scissors' problem treat the inter-sectoral terms of trade as an exogenous policy tool, other studies treat it as an endogenous variable and estimate it for a specific country. For example, Bilginsoy (1997) provides and tests a two-sector model of endogenous determination of the terms of trade for Turkey.

of agricultral goods dominated any negative impact of the required higher wages in the industrial sector to absorb the higher agricultural output, vice versa.[10] The impact of *KI(-1)* would be unambiguously positive, as the higher production capacity in the industrial sector increases the investible surplus that the state can extract in the form of industrial goods.The impact of *NE* could be positive or negative. It would be positive if the excess supply of agricultural products were exported and the resulting higher income of the rural/agricultural sector were extracted away by the state, as argued descriptively earlier by Preobrazhensky (1926, translated1965) and highlighted analytically later by Carter (1986) and Baland (1993). It would be negative if the foreign-credit financed imports of agricultural and industrial goods helped the state increase its accumulation at least in the short run, as suggested by Advocate (2008).

The reduced-form multivariate relationships derived from the Sah-Stiglitz's theoretical models of the 'price scissors' problem are generally estimated using traditional econometrics methods a' la ordinary least squares (OLS), generalized least squares (GLS), instrumental variables (IV), etc. (Li-Tsui (1985, 1990), Advocate (2008)). However, the reliability of these studies are undermined by the statistical insignificance of their estimated coefficients for certain intrinsic determinants of the investible surplus such as the agricultural and industrial production capacities as well as by the failing of their general test statistics such as the Durbin-Watson statistic (DW).[11] Their first

[10] This is for an economy closed, at least at the margin. Otherwise, it would be positive categorically if trade were possible and any excess supply of agricultural goods were exported instead of raising the urban/industrial wages to absorb it domestically. For details, see Carter (1986) and Baland (1993).

[11] Li-Tsui (1990) finds a statistically insignificant impact for both the agricultural and industrial capital-labor ratios in all their five regressions and low DWs for their OLS and IV estimations. Advocate (2008) finds a statistically insignificant impact for the agricultural capital stock per capita in two of her four regressions, a statistically significant negative impact for it in her other two regressions, and low DWs for three of her four OLS estimations.

estimation shortcoming reflects high correlation (multicollinearity) among their explanatory variables, while their second one indicates serial correlation (autocorrelation) in the error terms of their OLS and IV regressions.[12]

The consequential multicollinearity issue of the abovementioned studies of the 'price scissors' problem arises from estimating what is in effect a market-driven relationship with data from a centrally planned economy. The reduced-form multivariate relationships estimated in these studies imply that *IS* is endogenous and its determinants *TT*, *KA(-1)*, *KI(-1)*, and *NE* are exogenous and independent of one another. Also, their estimated relationships imply a unidirectional causation from the changes in *TT*, *KA*, *KI*, and *NE* to the changes in *IS*. While these requirements can be met in the invisible-hands world of a market economy, they are bound to be violated in the command world of a centrally planned economy where the target values of such aggregate variables are set in relation to one another and no one variable is given an absolute priority over the others for long. In addition, to the extent that the original and extended versions of the Sah-Stiglitz's models depend on market equilibrium conditions, the reduced-form multivariate relationships derived from them become further inappropriate for testing the validity of P1 in centrally planned economies where consumer rationing is standard practice.[13] As for the autocorrelation issue of the abovementioned studies, it arises from applying non-time series estimation methods such as OLS and IV to non-stationary time series data with inherent time trends.

[12] Li-Tsui (1990) correct for the low DWs but do not address why the agricultural and industrial production capacities play no role in their study. Advocate (2008) notes the possibility of multicollinearity between the agricultural and industrial capital stock per capita as well as the existence of autocorrelation in the error terms but does not correct for these problems.
[13] For an alternative modeling with consumer rationing, see Baland (1993).

This paper proposes and uses an alternative bivariate, as opposed to a multivariate, approach along with time-series estimation methods that circumvents the inherent multicollinearity and autocorrelation shortcomings of the aforementioned studies of the validity of P1 in a centrally planned economy. It shifts the focus of the analysis to stochastic bivariate specifications involving various pairs of the five variables of interest: *IS, TT*, *KA*, *KI*, and *NE*. This approach allows for examining not only the relationship between the investible surplus per capita and each of its assumed four determinants, but also the relationships among those four determinants themselves.

As for the time-series estimation methods, three sets of tests are performed. First, the stationarity and the order of integration of the five variables are examined using the augmented Dicky-Fuller test (ADF) for unit roots. The existence of unit roots in these variables would indicate non-stationarity in them, rendering unreliable a traditional specification and estimation of a relationship among them a' la OLS, GLS, IV, etc.[14] Second, the long-run relationships for various pairs of the five variables are studied using the Johansen test for cointegration. Of interest is not only the long-run relationship between the investible surplus per capita, *IS*, and each of its assumed four determinants (*TT, KA, KI,* and *NE*), but also the long-run relationship for various pairs of the four 'so-called' independent variables. The existence of pair-wise cointegration among the assumed determinants of *IS* would indicate an inherent multicollinearity among them, rendering a simultaneous inclusion of them in a multivariate specification inadmissible. Third, the short-run relationships a' la causality for various pairs of the five variables are examined using a Granger causality Wald test. As with the cointegration tests, of interest

[14] GLS partially corrects for non-stationarity of the variables by partial differencing of them but not for their multicollinearity.

is not only the direction of causation between the changes in the investible surplus per capita, ΔIS, and the changes in each of its assumed four determinants (ΔTT, ΔKA, ΔKI, and ΔNE), but also the direction of causation for various pairs of the latter four variables. The existence of pair-wise bidirectional causation between the changes in investible surplus per capita and the changes in each of its assumed four determinants renders unreliable the commonly-used multivariate specification and estimation of a relationship among them. The existence of pair-wise bidirectional causation among the assumed determinants of *IS* would indicate an inherent multicollinearity among them, rendering a simultaneous inclusion of them in a multivariate specification inadmissible.

As for the central hypothesis tested, in the multivariate framework of the existing studies with traditional estimation methods the validity of the Preobrazhensky's First Proposition, P1, requires only a statistically significant negative coefficient for the terms of trade, *TT*. In the bivariate framework of this study with its times-series estimation methods the validity of P1 requires not only a statistically significant negative cointegrating relationship between *IS* and *TT*, but also a unidirectional causation from ΔTT to ΔIS at least in the short run.

As for the data used, they are from the databank at the Lodz University in Poland and the Polish Main Statistical Data Office publications, same as in Advocate (2008). They are annual observations for the period 1960-1987. Their logarithmic transformation is used in this study, except for net exports per capita, *NE*, which can and does assume negative values for most years in the sample. This reinforces the transformation of the data into stationary series. The data for the two sectoral capital stocks are compiled from actual data reported in the above sources. The original division of 'urban vs. rural' in the ‘price scissors’ analysis is used for computing per capita values of various variables. The urban population is used for converting into per capita the investible surplus, the urban/industrial capital stock, and the net exports; the rural population is used for converting into per capita, the rural/agricultural capital stock.

3. The Relationship between the Investible Surplus and its Assumed Determinants

This section tests the validity of the Preobrazhensky’s First Proposition, P1, for Poland during the period 1960-87 using the bivariate approach proposed in Section 2 along with stationarity, cointegration, and Granger causality analyses. The specifics of the concepts of stationarity, cointegration, and Granger causality used in this study along with the respective estimation results are as follows.

3.1 Unit Roots Analysis

Testing for stationarity of variables of interest in a study does not provide any information about the relationships between them but satisfies the preconditions for studying their relationships. It also indicates whether traditional methods of estimations such as OLS and IV are appropriate for capturing their relationships or not. To these ends, the stationarity and the order of integration of the five time series *IS*, *TT*, *KA*, *KI*, and *NE* is tested using the augmented Dicky-Fuller test (ADF) for unit root. Using the investible surplus per capita, *IS*, as an example and following the standard notations in econometrics literature, this test is applied by positing the equation

$$\Delta IS(t)=\mu+\beta.t+\gamma^{*}IS\,(t-1)+\sum \text{Ø}_j\,\Delta IS(t-j)+\varepsilon(t)$$

where μ is the drift term, β is the coefficient of the deterministic time trend t, $\gamma^{*}= (\sum\gamma_i)-1$ for $i=1$ *to* p, $\text{Ø}_j=-\sum\gamma_k$ for $k=j+1$ *to* p, Δ is the difference operator, p is the order of autoregressive process (AR) and $\varepsilon(t)$ are white noise. The unit root test for this model is carried out by testing the joint hypothesis that $\beta = \gamma^{*}=0$. In the absence of a time trend the unit root test for the model is carried out by testing the hypothesis that $\gamma^{*}=0$.

Since unit root tests are sensitive to the inclusion/exclusion of an intercept and a time trend as well as to the number of included lags, alternative models are considered and the one that works for all variables is selected. The AR(1) model with no intercept and no time trend, that is the pure random walk model, does not work for *KA* and *KI*. The AR(1) model with an intercept and a time trend works for all variables except *IS*. The standard AR(1) model with intercept works for all variables and, hence, is selected.

The test statistic and P-value found by applying the ADF unit root procedure to the five time series are presented in Table 1. The null hypothesis is the presence of unit root and a larger negative test statistic is a rejection of this hypothesis. All five time series are integrated of order 1 and non-stationary, and their first differences are integrated of order 0 and stationary.[15] This renders inappropriate the use of traditional estimation methods such as OLS and IV for estimating any relationships among the five non-stationary variables of interest, as hypothesized in Section 2. Such incorrect application of estimation methods would manifest itself via autocorrelation in the error terms of the estimated regressions, as in the studies discussed in Section 2. The testing for stationarity and the order of integration of the variables just completed satisfies the precondition for studying the short- and long-run relationships between them next.

[15] Taking logs, first differences, or both usually transforms non-stationary data into stationary series. Here the logarithmic transformation of the five series did not make them stationary, but helped making *IS* stationary with first differencing.

3.2 Cointegration Analysis

The long-run relationships for the various pairs of the five time series are analyzed using the Johansen test for cointegration that is based on Engle and Granger's approach to analyzing cointegration. The Johansen method is a full information maximum likelihood estimation of a system of cointegrating relationships based on the VAR approach. This method can be expressed for the bivariate approach of this paper as follows. Let X_t be a *2 x 1* vector of *2* stochastic variables that are integrated of order 1. Then X_t can be written as a pth order VAR that can be represented in the error correction form (ECM)

$$\Delta X_t = \Pi X_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta X_{t-i} + \Psi D_t + \varepsilon_t$$

where Δ is the difference operator, $\Pi = (\sum_1^p \Pi_i - I)$,[16] $\Gamma_i = -\sum_{j=i+1}^{p} \Pi_j$, and D_t is a *d x 1* vector of deterministic terms, typically a 1 to capture the constant in each equation, the time trend *t*, and 'intervention' dummy variables, as needed. Ψ is the *2 x d* matrix of coefficients associated with D_t. Finally, the vector of error terms ε_t is *2* dimensional zero-mean random variables with variance matrix Ω.

The appeal of the above ECM is its explicit distinction between long-run equilibrium and dynamic adjustments to it.[17] Its transparent display of the long-run cointegrating relationship between the two variables in X_t is of interest here. To that end, if there are *r (r<2)* independent linear combinations of X_t which are difference stationary, then X_t is cointegrated of order *r*. According to the Granger representation theorem, if

[16] $\Pi_i (i = 1,..., p)$ is the coefficient of X_{t-i} in the original pth order VAR.

[17] For details, see Patterson (2000), section 14.4.

X_t is cointegrated of order *r*, then the *2 x 2* matrix Π has rank *r* (*r<2*) and one can write $\Pi = \alpha\beta'$, where both α and β are *2 x r* matrices of full column rank.[18] The Johansen method is basically a procedure for estimating the above relationship subject to the constraint $\Pi = \alpha\beta'$.

The order of the VAR, *p*, is determined in advance by lag selection criteria such as AIC (Akaike Information Criterion), HQIC (Hannan-Quinn Information Criterion), and SIC/BIC/SBIC (Schwarz Bayesian Information Criterion).[19] Considering the orders of 1 and 2, HQIC and SBIC suggest a lag length of 1 and AIC suggests a lag length of 2 for the five variables (*IS, TT, KA, KI, NE*) as a group. Given this and the likely presence of a time trend in the time series data, the model with a lag length of 1 and a constant and a time trend is the default one used. Where it does not provide a conclusive result for a particular pair, it is modified at the margin, for the lag length and/or exclusion of a time trend. The stability of the results is tested by considering alternative models, even with lag orders higher than 2.

The pair-wise cointegration results for the five series are presented in Table 2.[20] The first column states the specific pair of variables whose long-run relationship is being tested. The second column lists the type of the deterministic terms and the lag length in the estimated model. The third column reports the trace test statistics,[21] which are then

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[18] That is, r is also the column dimension of α and β. α contains the adjustment coefficients and β contains the equilibrium/cointegrating coefficients allowing for separate representation of the two coefficient sets. β' is the cointegrating vector.

[19] It should be noted that asymptotically, as the sample size approaches infinity, AIC overestimates the true order of the autoregression (the true lag length for the AR) and is said to be inconsistent, whereas HQIC and SIC do not and are said to be consistent. For details, see Patterson (2000).

[20] The number of observations in these estimations is mostly (8 out of 10) 25, with the lowest being 24 for the pair *TT-KI(-1)* with a lagged variable and a 2nd order VAR model. The results were obtained using Stata.

[21] Johansen develops two test statistics for determining the cointegration rank. The first test is known as the

compared with the critical values in the fourth column. A trace statistic larger than the critical value provides evidence against the null hypothesis of *r* or fewer cointegrating vectors. The fifth column indicates the number of cointegrating relationships hypothesized. The first hypothesis ($H_0: \; r = 0$) tests whether the cointegration rank is zero--there is no equilibrium condition that keeps the considered pair of variables in proportion to each other in the long run. This hypothesis is rejected for all pairs of variables in this study except the ones that involve net exports per capita, *NE*. The second hypothesis ($H_0 : r \leq 1$) tests whether the cointegration rank is less than or equal to one for all pairs of variables that do not involve *NE*. The Johansen test failed to reject this hypothesis for any of the six pairs of variables under consideration. This combined with the results of the first hypothesis brings us to the conclusion that the cointegration ranks for all the pairs of variables that exclude *NE* equals one. The normalized estimated cointegrating coefficients for each of these pairs of variables and their respective t-statistics are reported in the last column of Table 2.

The cointegration analysis finds a long-run relationship for all pairs of variables that do not involve *NE*.[22] Most importantly, it finds a significant negative relationship between *IS* (investible surplus per capita) and *TT* (relative price of agricultural to industrial goods) in line with the theoretical result in Sah-Stiglitz (1992). Given the central role of this negative relationship in the 'price-scissors' economics, the robustness

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trace statistics and is the relevant test statistics for the null hypothesis $r \leq r_0$ against the alternative $r \geq r_0 + 1$. The second test is the maximum eigenvalue test and improves the power of the test by changing the alternative hypothesis to $r = r_0 + 1$. The first test is used in this study. For details, see Patterson (2000), section 14.4.3.

[22] Similar results emerge if *KA* and *KI* are used instead of their lagged values. For comparison and symmetry with the existing empirical studies of 'price-scissors' problem, the results for the lagged *KA* and *KI* are reported.

of this finding is tested by considering different model specifications for the pair *IS-TT*. The cointegrating rank for this pair is 1 for the lag length of 1 and 2 and for the model with and without a time trend, and three of the four estimated beta coefficients are negative and statistically significant.[23] Thus, it is concluded that there was a significant negative long-run relationship between *IS* and *TT* in Poland for the period 1960-1987.

As for the other pairs of variables, the cointegration analysis finds a significant positive relationship between *IS* and *KA(-1)* (the lagged capital stock per capita in the agricultural sector) as well as between *IS* and *KI(-1)* (the lagged capital stock per capita in the industrial sector). These positive relationships are in line with the theoretical expectation that over time an increase in the capital stock per labor (capita) in each sector would be associated with a rise in production and, hence, in the investible surplus per capita. As the word "associated" implies, the previous statement can be stated in reverse order and the causality can run either way, as discussed in the next section. The next two cointegration tests reveal a significant negative relationship between the relative price of the agricultural to industrial goods, *TT*, and each of the two sectoral capital stocks per capita, *KA(-1)* and *KI(-1)*. This result can be expected by transitivity, given the previous cointegration results regarding the pairs *IS-TT*, *IS-KA(-1)* and *IS-KI(-1)*.[24] By the same reason, the next cointegration result that is the significant positive relationship between the lagged capital stock per capita in the two sectors, *KA(-1)* and *KI(-1)*, can be expected. Interestingly, their normalized cointegrating coefficient (Beta) of +1.04 indicates that

[23] Using a lag length higher than 2 seems to place us beyond the test of r=0 or r=1 with no clear result. For symmetry with the model used for other pairs, the case with the constant, time trend and lag length 2 is reported in Table 2.
[24] When higher *IS* is associated with lower *TT* and higher *IS* is associated with higher *KA(-1)* and *KI(-1)*, then lower *TT* would be associated with higher *KA(-1)* and *KI(-1)*, vice versa.

they were changed by the central planners in unison and the capital stock per capita in the industrial sector was on average only 4% higher than that in the agricultural sector.[25]

The pair-wise cointegration among *TT, KA(-1),* and *KI(-1)* renders simultaneous inclusion of them in any multivariate analysis of *IS* inadmissible. Their close link results in multicollinearity and, hence, in incorrect estimation of their individual impact on *IS*, as discussed in Section 2.

The last four pair-wise cointegration tests reported in Table 2 consider the much-discussed role of trade for industrialization. In theory exports of the excess supply of agricultural goods can entail a higher domestic investible surplus as well as a higher imports of capital equipments and, hence, can foster industrialization over time. Also, access to international credit can do the same. However, the cointegration analysis finds no long-run relationship for any pair that involves *NE* for the models considered.[26] The robustness of this result is mostly confirmed by different model specifications as follows. The cointegrating rank for the pair *IS-NE* is zero (r=0) for the lag length of 1 and 2 and for the model with and without a constant and a time trend. This result holds even with higher lag lengths for the model with a constant and a time trend. The cointegrating rank for the pair *TT-NE* is zero (r=0) for the lag length of 1 to 4 for the model with a constant. It is zero also for the model with a constant and a time trend, but only for the lag length of 2. Using a lag length different from 2 along with a time trend places us beyond the test of r=0 or r=1 with no clear result. For symmetry with the models used for other pairs, the case with lag length 2 with a time trend is reported in Table 2. For the pairs *KA(-1)-NE* and *KI(-1)-NE*, their test results from the cointegration rank and the Beta identification

[25] Naturally, the same exact results hold for the non-lagged values of the two sectoral capital stocks per capita variables.
[26] The same no cointegration results emerge for exports and imports separately as well.

are in conflict when a time trend is included and higher lags are considered. Their cointegrating rank is clearly zero (r=0) for the model with a constant and a lag length of 1, as reported in Table 2.

These results of no cointegration (r=0) for the pairs that involve *NE* are corroborated by the following facts. During this period the Polish foreign trade was constrained, for example, by weak external demand for its exports and by its limited access to international credit. Also, the trade deficits that Poland ran each and every year between 1960 and 1981 were mostly small (less than 5%) relative to its total investment outlays.[27] Thus, it is concluded that trade did not play any measurable long-run role in the total (capital) accumulation, the sectoral capital deepening, or the determination of domestic inter-sectoral terms of trade in Poland during the period 1960-87.

As Maddala and Kim (1999) point out, when two variables are cointegrated, at least one must Granger cause the other. Also, there can be a Granger causal relationship between the transformed stationary values of two non-stationary variables that are not cointegrated. Thus all of the ten pairs of variables listed in Table 2 are tested for causality as follows.

3.3 Granger Causality Analysis

Theoretically in the invisible hands world of a market economy, where *IS* is determined endogenously, the direction of causation would be unidirectional from changes in *TT*, *KA*, *KI*, and *NE* to changes in *IS*. But in the command world of a centrally planned economy, where target values for *IS* and other variables are set for each period exogenously, the direction of causation depends on the central planners' priorities and

[27] In the early 1970s the government made an effort to increase the net imports, and for six of the years between 1973 and 1981 the trade deficits relative to total investments were over 10%, topping at 15% in 1976.

"policy rules". If the central planners give an absolute/unconditional priority to *IS* for each and every period, then changes in *IS* lead (Granger cause) changes in other variables. In this case, for example, *TT* is changed as needed to make the target change in *IS* feasible. On the other hand, if the central planners give an absolute/unconditional priority to *TT* or other non-*IS* variables, then changes in *TT* or other non-*IS* variables lead (Granger cause) changes in *IS*. A third possibility of bi-directional causality arises if the central planners shift back and forth their priorities across various variables so as to keep different parts of the economy in certain proportion with one another. This would ensue, for example, when the central planners follow a set of balanced growth and development "policy rules".

The causality for various pairs of the five time series *IS, TT, KA, KI,* and *NE* is examined using the Granger causality Wald test. Using the investible surplus per capita, *IS*, and the relative price of agricultural to industrial goods or the inter-sectoral terms of trade, *TT*, as an example and following the standard notations in econometrics literature, this test is applied by positing the following closed *p*th-order bivariate vector autoregressive system, the VAR

$$\Delta IS_t = \mu_0 + \sum_{k=1}^{p} \mu_{1k} \Delta IS_{t-k} + \sum_{k=1}^{p} \mu_{2k} \Delta TT_{t-k} + u_{1t}$$

$$\Delta TT_{\mathrm{t}} = \theta_0 + \sum_{k=1}^{p} \theta_{1k} \Delta TT_{t-k} + \sum_{k=1}^{p} \theta_{2k} \Delta IS_{t-k} + u_{2t}$$

where ΔIS and ΔTT represent the transformed stationary values of the variables *IS* and *TT* respectively, μ_0 and θ_0 are the drift/constant terms, p is the predetermined order of the VAR, and u_{1t} and u_{2t} are the standard (white noise) error terms. If the lags of ΔTT can improve a forecast for ΔIS in the presence of the lags of ΔIS, then ΔTT is said to Granger cause ΔIS. More specifically, in the first VAR equation, under the null hypothesis that ΔTT does not Granger cause ΔIS, then all of the μ_{2k} coefficients would be statistically equivalent to zero. Similarly, in the second VAR equation, under the null hypothesis that ΔIS does not Granger cause ΔTT, then all of the θ_{2k} coefficients would be statistically equivalent to zero

The order of the VAR, or the length of the longest lag in the autoregression, can be determined by lag selection criteria such as AIC, HQIC, and SIC/BIC/SBIC, as noted in the previous section on cointegration. All three information criteria suggest a lag length of 4 for the five stationary variables ΔIS, ΔTT, ΔKA, ΔKI, and ΔNE as a group. This lag order of 4 is used for their pair-wise VARs for consistency across the causality tests for various pairs of variables. The lag orders lower and higher than 4 are used to test the stability of the results.

The causality test results in Table 3 reveal that at the predetermined VAR order of 4 the causation is bi-directional for six of the ten pair-wise combinations of the five variables. Even for the remaining four pairs that have unidirectional causation, the causation becomes bidirectional when the VAR order is raised above 4.[28] These results indicate that the central planners did not give an absolute priority to *IS* or any of the other four variables *TT, KA, KI,* and *NE* for the entire period. They seem to have adjusted their target values for these variables in relation to one another to keep a balance among them over time.

The specific implications and stability of the results reported in Table 3 can be analyzed by considering the VAR orders below and above 4 as follows. Regarding the

[28] The small size of the sample limits the upper bound for the order of the VAR. At the order of 6 the difference between the number of observations and that of estimated parameters falls below ten.

pair *IS*-*TT* that has the central role in the ‘price scissors’ problem, when the order of the VAR is set below 4 a strong causation from the changes in *IS* to the changes in *TT* along with a weak one from the changes in *TT* to the changes in *IS* emerge at the order of 3.[29] When the order of the VAR is set above 4 the strong bidirectional causation between the changes in these two variables established at the order of 4 persists. This rejects the principal assumption of the existing studies of the ‘price scissors’ problem that the central planners give an absolute priority to the investible surplus and simply set the inter-sectoral terms of trade between agriculture and industry to achieve their target *IS*. In Poland they seem to have adjusted their target values for these two key variables in relation to one another, perhaps in response to opposite political pressures from the urban and rural populations.

For the pair *IS*-*KA*, there is a persistent evidence for bidirectional causality at all orders of the VAR except the order of 1 when there is a unidirectional causality from the changes in *KA* to the changes in *IS*. This suggests that the central planners adjusted their target values for *IS* and *KA* in relation to each other even on a short two-year time horizon, in line with their intention to extract the most from the agricultural sector.

Regarding the pair *IS*-*KI*, when the order of the VAR is set at 3 as well as above 4 the same bidirectional causation between the changes in the two variables listed in Table 3 recurs. However, when the order of the VAR is set at 1 and 2 a weak (at 10.5%) and a strong (at 0.8%) unidirectional causality from the changes in *KI* to the changes in *IS* appears, respectively.[30] This indicates that the changes in the industrial capital stock per

[29] There is no Granger causation between these two variables at the VAR orders of 1 and 2.
[30] The unidirectional causality from the changes in *KA* and *KI* to those in *IS* at the low VAR orders of 1 and 2 mirrors the relationship between the production capacity in the two sectors and the investible surplus per capita presented by the reduced-form multivariate relationships derived from the Sah-Stiglitz’s

capita leads (Granger causes) the changes in the investible surplus per capita for one to two years but after that the causation runs both ways. That is, the central planners' focus on the stock of industrial capital and industrialization as such did not dominate their other concerns such as the investible surplus continuously.

The causality between the changes in the relative price of agricultural to industrial goods, ΔTT, with the changes in each of the two sectoral capital stocks per capita, ΔKA and ΔKI, are as follows. For the pair *TT-KA*, when the VAR order is set below 4 or above 4 the causation remains unidirectional from the changes in the agricultural capital stock per capita to the changes in the terms of trade until the VAR order of 6 where it becomes bidirectional.[31] Similar results hold for the pair *TT-KI* with one difference, the causation becomes bidirectional one lag later, at the VAR order of 7. These results suggest that the central planners changed the relative price of agricultural to industrial goods in relation to the (previous) changes in the two sectoral capital stocks per capita. In effect, they changed the relative price of the two goods in accordance with the changes in the production capacity of the two sectors, and not by some abstract rule.

For the pair *KA-KI*, when the order of the VAR is set below as well as above 4 the causation remains bidirectional. This result indicates that the central planners changed the capital stock per capita in the two sectors in unison, as also established by the long-run cointegrating relationship between them discussed in the previous section. The short- and long-run tight relationships between the two sectoral capital stocks per capita are a main reason why multivariate specifications and estimations involving both variables are

theoretical models of the 'price scissors' problem.

[31] There is no Granger causation between these two variables at the VAR order of 1.

rendered inappropriate for studying ‘price scissors’ in Poland, and perhaps in other centrally planned economies.

The last eight tests reported in Table 3 examine in effect the short-run role of trade. Regarding the pair *IS-NE*, when the order of the VAR is set below 4 the causation remains unidirectional from the changes in *IS* to the changes in *NE* and when it is set above 4 the causation becomes bidirectional. This result confirms the idea that the central planners used trade (the net imports) to reach their target *IS* for short periods of few years. But they could not treat the trade balance residually beyond few years because of the well-known international credit and trade restrictions that they faced.

For the pair *TT-NE*, when the order of the VAR is set below 4 the causation remains unidirectional from the changes in *TT* to the changes in *NE* and when it is set above 4 the causation becomes bidirectional. This result indicates once again that the central planners used trade to accommodate their internal targets, here domestic terms of trade, for a few years but ultimately had to give consideration to the trade balance as well.

Regarding the pair *KI-NE*, when the order of the VAR is set below 4 (at 2 or 3) the causation becomes unidirectional from *NE* to *KI* and when it is set above 4 the causation remains bidirectional.[32] For the pair *KA-NE*, changing the VAR order generates similar results with one difference: the causation becomes bidirectional sooner at the VAR order of 3 rather than 4. These results indicate that the central planners adjusted the capital stock per capita in agriculture and industry in response to the changes in the trade balance in the previous 2 to 3 years. They seem to have adjusted the production capacity

[32] There is no Granger causation between these two variables at the VAR order of 1.

in the two sectors to control their trade balance and, hence, to deal with their trade restrictions.

The bidirectional causality between various pairs of the five variables at the high orders of the VAR, meaning over periods of three, four or more years, may not be surprising given the interdependence among such aggregate variables. But, lack of a unidirectional causality at the VAR orders of 1 and 2, meaning over periods of one and two years, in certain pairs, especially the investible surplus per capita and the inter-sectoral terms of trade between agriculture and industry is significant. This finding rejects the fundamental assumption in the 'price scissors' economics that the terms of trade is merely an exogenous policy tool used to obtain target values of the investible surplus, at least in the short run.

4. Conclusions

This paper has studied the validity of the Preobrazhensky's First Proposition, P1, for the centrally planned Poland during the period of 1960-1987 by testing whether the state increased its internal accumulation, investible surplus, by setting the inter-sectoral terms of trade against its agriculture, in favor of the industry. It has revealed that the existing multivariate empirical approach derived from the Sah-Stiglitz's market-based theoretical model of the 'price scissors' problem can misrepresent not only the validity of P1 in a centrally planned economy but also the impact of the intrinsic determinants of the state's investible surplus such as the production capacities in agriculture and industry. It has used an alternative bivariate approach along with time-series estimation methods of cointegration and Granger causality analysis to circumvent the inherent multicollinearity and autocorrolation shortcomings of the existing studies.

As for the central hypothesis tested in this study, the significant negative long-run cointegrating relationship between the investible surplus per capita and the relative price of agricultural to industrial goods seems to support the validity of the Preobrazhensky's First Proposition in Poland for the period 1960-1987. However, the lack of a unidirectional causality from the changes in the inter-sectoral terms of trade to the changes in the investible surplus per capita even for short periods of one to two years rejects the fundamental assumption of this proposition that the inter-sectoral terms of trade is an exogenous policy tool used to obtain target values of the investible surplus, at least in the short run.

As for the two intrinsic determinants of the internal accumulation, there is a significant positive long-run cointegrating relationship between the investible surplus per capita and the capital stock per capita in the agricultural and industrial sector each. There is also a unidirectional Granger causality from the changes in the capital stock per capita in the two sectors to those in the investible surplus per capita over short periods of one to two years. These results reflect the innate role of the sectoral production capacities in determining the internal accumulation, a role that can be missed by the multivariate approach of the existing studies.

The bivariate approach used in this study has allowed for examining not only the relationship between the investible surplus per capita and each of its assumed four determinants, but also the relationships among those four determinants themselves, as summarized next. There is a significant negative long-run cointegrating relationship between the relative price of the agricultural to industrial goods and each of the two sectoral capital stock per capita. Also, there is pair-wise unidirectional Granger causality from the changes in the the two sectoral capital stock per capita to the changes in the inter-sectoral terms of trade, suggesting that the central planners changed the relative price of agricultural to industrial goods in accordance with the changes in the production capacity of the two sectors, and not by some abstract rule. Regrading the two sectoral production capacities themselves, there is not only a significant positive long-run cointegrating relationship between the levels of the capital stock per capita in the two sectors, but also a near unity normalized cointegrating coefficient between them. In addition, there is an unequivocal bidirectional Granger causality between the changes in them. These results indicate that they were set and changed by the central planners in unison. The tight pair-wise relationships among the terms of trade and the two sectoral capital stock per capita renders inappropriate any multivariate specification and estimation that uses these three variables for studying the 'price scissors' problem in Poland, and perhaps in other centrally planned economies.

As for the much-discussed role of trade and external funding for industrialization, while the cointegration analysis did not find any long-run relationship between the net exports per capita and any of the other four variables, the Granger causality analysis found short-run relationships between the changes in the net exports per capita and the changes in the other four variables. These results indicate that trade could and did play a role in targeting and managing main aggregate variables such as the internal accumulation, the sectoral capital stocks, and the inter-sectoral terms of trade in the centrally planned Poland during the period 1960-1987, but only in the short run because of the well-known international credit and trade restrictions that the country faced.

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Table 1: Unit Root Tests

Variable	Test Statistic	P-value
IS	-1.761	0.4001
ΔIS	-2.633*	0.0865
TT	-1.736	0.4129
ΔTT	-6.355***	0.0000
KA	-1.716	0.4230
ΔKA	-11.313***	0.0000
KI	-1.567	0.5003
ΔKI	-12.574***	0.0000
NE	-1.188	0.6788
ΔNE	-4.049***	0.0012

Notes: The listed MacKinnon P-values are based on the interpolated Dickey-Fuller critical values of CV(10%)=-2.63, CV(5%)=-3.00, and CV(1%)=-3.75. A larger negative test statistic than these critical values is a rejection of the null hypothesis of the presence of unit root at 90% (*), 95%(**), or 99%(***) confidence level respectively. The standard AR(1) model with intercept is used for all variables; and Δ is the difference operator. The number of observations is 27 and 26 for the level and differenced form of the variables.

Table 2: Pair-wise Cointegration Tests

Variable Pair	Model[a]	Trace Statistic[b]	CV(5%)	H_0	Normalized Beta of r=1[c]
IS – TT	C, T, 2 lags	19.17 3.34**	18.17 3.74	r = 0 r ≤ 1	-9.55 (0.001)
IS – KA(-1)	C, T, 1 lag	27.35 1.60**	18.17 3.74	r = 0 r ≤ 1	+7.81 (0.000)
IS – KI(-1)	C, T, 1 lag	30.83 1.31**	18.17 3.74	r = 0 r ≤ 1	+3.83 (0.000)
TT – KA(-1)	C, T, 1 lag	20.05 2.81**	18.17 3.74	r = 0 r ≤ 1	-2.05 (0.000)
TT – KI(-1)	C, 2 lags	26.14 0.77**	15.41 3.76	r = 0 r ≤ 1	-0.47 (0.000)
KA(-1)– KI(-1)	C, T, 1 lag	18.89 0.97**	18.17 3.74	r = 0 r ≤ 1	+1.04 (0.006)
IS – NE	C, T, 1 lag	16.70**	18.17	r = 0	NA
TT – NE	C, T, 2 lags	17.11**	18.17	r = 0	NA
KA(-1) – NE	C, 1 lag	7.07**	15.41	r = 0	NA
KI (-1) – NE	C, 1 lag	6.49**	15.41	r = 0	NA

Notes: [a] C and T denote constant and time trend. [b] A trace statistic larger than CV(5%) is a rejection of the null hypothesis of the cointegration rank being less than or equal to r (r=0, 1) at 95%(**) confidence level. [c] The cointegrating coefficients (Beta) for the case r=1 presented are normalized on the first variable in each pair with their t-statistics in parentheses. The number of observations for the fifth and seventh estimations are 24 and 26 respectively, and 25 for the other eight estimations.

Table 3: Pair-wise Granger Causality Wald Tests

Equation	Excluded	Chi2 Statistic	df	Prob > Chi2
ΔIS	ΔTT	12.068	4	0.017
ΔTT	ΔIS	15.150	4	0.004
ΔIS	ΔKA	9.496	4	0.050
ΔKA	ΔIS	33.551	4	0.000
ΔIS	ΔKI	9.248	4	0.055
ΔKI	ΔIS	8.595	4	0.072
ΔTT	ΔKA	12.499	4	0.014
ΔKA	ΔTT	5.146	4	0.273
ΔTT	ΔKI	15.905	4	0.003
ΔKI	ΔTT	2.572	4	0.642
ΔKA	ΔKI	16.957	4	0.002
ΔKI	ΔKA	44.193	4	0.000
ΔIS	ΔNE	6.181	4	0.186
ΔNE	ΔIS	12.113	4	0.017
ΔTT	ΔNE	3.968	4	0.410
ΔNE	ΔTT	10.672	4	0.031

Continued…

Table 3: Pair-wise Granger Causality Wald Tests (Continued)

Equation	Excluded	Chi2 Statistic	df	Prob > Chi2
ΔKA	ΔNE	25.342	4	0.000
ΔNE	ΔKA	14.473	4	0.006
ΔKI	ΔNE	13.299	4	0.010
ΔNE	ΔKI	11.055	4	0.026

Notes: The listed probability values, Prob, are based on right (as opposed to left) critical values for the Chi2 distribution. A Chi2 statistic larger than the respective critical values is a rejection of the null hypothesis that the coefficients on the lags of the variable in the "excluded" column are jointly zero in the VAR equation for the variable in the "equation" column, at the (1-Prob)% confidence level. The degrees of freedom of the test, "df", is the number of zero restrictions on the coefficients of the lags of the variable in the "excluded" column and reflects the order of the VAR estimated. The standard VAR with intercept is used for all tests; and Δ is the difference operator. The number of observations for all tests is 23.