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Article

Econometric Assessment of Financial and Economic Factors Driving Construction Project Cost Growth

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Abstract: The article uses econometric methods to assess the financial and economic factors behind construction cost growth. First, using monthly data for 2021–2026 (n = 66), construction and installation prices are regressed on consumer prices, material prices, wages and import prices with Newey–West standard errors (R^2 = 0.903). Second, using panel data for 14 regions in 2012–2022 (n = 154), fixed- and random-effects models estimate the effect of financing on real output and construction costs. Materials account for 41.3%, the exchange rate for 29.7% and wages for 22.4% of cost growth, while 27.9% of additional financing is absorbed by price increases.

Keywords: econometric model, multiple regression, panel data, construction cost, financing efficiency, elasticity, Driscoll–Kraay, price loss coefficient.

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1. Introduction

Construction cost growth directly affects the efficiency of centralized capital investment: the real volume of output that can be created with a given nominal amount of funds declines in proportion to price growth. In economic theory, the economic return on public capital investment is measured by the volume of physical capital it creates [1]. According to IMF estimates, the average country loses about 30 percent of the potential gains from public investment due to inefficiencies in its management. Cost overruns relative to plans in infrastructure projects have also been shown to be systematic rather than random. W. Baumol explained why prices rise faster than average inflation in sectors with slow productivity growth. The methodology for estimating construction cost escalation through resource indices is set out in the recommended practices of AACE International [2]. In time series, autocorrelation and heteroskedasticity are usually handled with Newey–West standard errors. For panel data with cross-sectional dependence, Driscoll–Kraay standard errors are used. The choice between fixed- and random-effects estimators is commonly based on the Hausman test. Cross-sectional correlation in the residuals is checked with the Pesaran test. A comprehensive treatment of these panel data methods is given by J. M. Wooldridge.

In Uzbekistan, I. O. Karimov analyzed the theoretical and practical foundations of investment activity in the construction sector [3]. He also studied foreign experience in managing investment in construction enterprises and ways of applying it in Uzbekistan. In a later study, he proposed approaches to assessing the factors that affect the

investment attractiveness of construction enterprises. I. K. G'iyosov examined how corporate governance affects the economic efficiency of construction enterprises [4]. He also identified priority directions for improving their governance mechanisms under digital transformation . Therefore, quantifying which financial and economic factors drive construction cost growth, and how strongly, is an important scientific and practical task for public investment planning. In practice, price changes are often attributed to headline inflation indicators. However, the actual drivers of construction cost, as well as the effect of financing itself on prices, have not been assessed econometrically using data for Uzbekistan [5].

The aim of the article is to assess, by econometric methods, the financial and economic factors driving the growth of construction project costs. The study was conducted in two stages. In the first stage, the effect of cost factors (inflation, material prices, wages and the exchange rate) was estimated with a multiple regression model on monthly data. In the second stage, the effect of the financing factor on real output and construction cost was estimated on regional panel data.

2. Materials and Methods

In this article, financing efficiency is understood as the degree to which funds allocated to construction and installation works are converted into real construction output (at constant prices). Construction cost growth affects this efficiency through two channels. The first is the direct (purchasing power) channel: when resource prices rise, the same nominal amount of funds yields a smaller real volume. The second is the demand (pressure) channel: a sharp increase in financing intensifies demand in the construction resource market and causes additional price growth, so that part of the additional funds is "absorbed" by price increases rather than translated into real output. International studies link the low efficiency of public investment precisely to cost growth at the implementation stage and show that cost overruns in construction projects are systematic. However, the magnitude of these effects has not been sufficiently quantified for developing countries, including on data for Uzbekistan.

Based on the above, three working hypotheses were put forward. H1: growth in real financing of construction and installation works has a positive effect on the growth of real construction output in the regions, but this effect is less than proportional (the elasticity is below one). H2: an increase in the national construction price index leads to an increase in regional construction costs, i.e., the price shock is transmitted to the regions. H3: growth in financing by itself raises regional construction costs (the demand channel); consequently, part of the additional financing is absorbed by price growth.

The theoretical basis of the models is a generalized construction production function. According to it, real construction output depends on the volume of real resources – materials, labor and machinery – which are purchased with monetary funds at current prices. If real output in region i is expressed as $Q_{it} = A_i \cdot (F_{it}/P_{it})^{\beta}$ (where F is nominal financing, P is the price of construction resources, and β is the elasticity with respect to real financing), then taking logarithms and first differences yields a linear equation in growth rates. In this equation, price growth affects real output with the same strength as financing growth but with the opposite sign. It is precisely this property that makes it possible to quantify the effect of price increases on financing efficiency: if nominal financing is unchanged, a 1 percent rise in prices reduces real output by β percent. Resource prices, in turn, are determined by overall price dynamics and by the demand for resources in the region. This provides the basis for the second equation – the cost equation.

To test the hypotheses, a panel dataset was compiled for the 14 regions of Uzbekistan (the Republic of Karakalpakstan, 12 provinces and Tashkent city). Regional data on the volume of construction works and their real growth rates were obtained from

the National Statistics Committee. Data on investment in fixed capital and the share of construction and installation works in it were drawn from the same source. The price index for investment in fixed capital and construction works was used to convert nominal values into real terms. The estimation period covers 2012–2022 (11 years × 14 regions = 154 observations). The lower bound of the period is determined by the need to compute the lagged value of financing, and the upper bound by the 2023 change in the methodology for calculating the volume of construction works (expanded coverage of the informal sector). This change created a break in the regional nominal series, and regional real growth rates for 2024 have not been published. All variables are expressed as logarithmic growth rates, so the coefficients are interpreted as elasticities.

Pairwise correlation and regression analysis estimates the effect of individual factors without separating them from one another. In practice, however, the cost of construction projects is shaped simultaneously by a number of factors. Therefore, at the next stage of the study, a multifactor econometric model of construction cost growth was developed. Theoretically, the change in construction cost can be expressed as the following function:

$$\Delta C = f(INF, MP, W, ER, T, F, X), \quad (1)$$

where ΔC is the growth in construction cost; INF is inflation; MP is the price of construction materials; W is wages; ER is the exchange rate; T is the construction period; F is the financing factor; X denotes other factors. The first four factors (INF, MP, W, ER) are construction cost factors that directly affect the estimated (budgeted) cost through resource prices. The construction period (T) determines the period over which price growth accumulates in the project cost. The financing factor (F) operates through the demand channel described above. Each factor of the function was operationalized using official statistical indicators (Table 1).

Table 1. Operationalization of the factors of the construction cost growth function

Symbol	Factor	Indicator used	Source, frequency	Expected effect
ΔC	Construction cost growth	Price index of construction and installation works	NSC, 1.11.02.0011, monthly	Dependent variable
INF	Inflation	Consumer price index (COICOP-2018)	NSC, 1.11.01.0002, monthly	+
MP	Price of construction materials	Price index of material costs in construction and installation works	NSC, 1.11.02.0011, monthly	+
W	Wages	Average monthly wage in construction	NSC, 2.02.02.1410, monthly	+
ER	Exchange rate	Import price index (in soum, 3-month average)	NSC, 1.11.02.0016, monthly	+
T	Construction period	Price escalation formula; lag in the panel model	Calculation; regional panel	+
F	Financing factor	Growth of investment in construction and installation works	NSC, regions, annual	+

X	Other factors	Intercept, regional effects, residual	–	–

Note: *NSC – National Statistics Committee of the Republic of Uzbekistan.*

Because the factors are observed at different frequencies, the model was estimated in two stages. In the first stage, the effect of the cost factors (INF, MP, W, ER) was estimated on monthly data from January 2021 to June 2026 (66 observations). Monthly frequency was chosen because the National Statistics Committee publishes the price index of construction and installation works and the consumer price index on a monthly basis. Wages in construction are also reported monthly. The same applies to the import price index. This increases the number of observations several-fold compared with annual data. In the second stage, the effect of financing (F) was estimated on regional panel data (panel models), and the effect of the construction period (T) was estimated through the lag structure of the panel model and the formula for the accumulation of prices in cost. Other factors (X) were accounted for by the model intercept, regional fixed effects and the random error term.

The first-stage model was specified in log-linear form:

$$\Delta lnC_t = \beta_0 + \beta_1 \cdot INF_t + \beta_2 \cdot MP_t + \beta_3 \cdot W_t + \beta_4 \cdot ER_t + \varepsilon_t, \qquad (2)$$

where $\Delta \ln C_t$ is the monthly growth of the price index of construction and installation works (relative to the previous month, log percent); INF_t is the monthly growth of the consumer price index; MP_t is the monthly growth of the price index of material costs (materials) in construction and installation works; W_t is the monthly equivalent of the growth of the average monthly wage in construction relative to the same month of the previous year (annual growth/12); ER_t is the three-month moving average of the monthly growth of the import price index. Import prices are expressed in soum and calculated from customs statistics; therefore, they reflect changes in the exchange rate and the pass-through of external prices to construction cost. Since all variables are measured in log percent, the β coefficients are interpreted as elasticities. The model parameters were estimated by ordinary least squares in matrix form:

$$b = (X'X)^{-1} \cdot X'y, \quad (3)$$

where y is the vector of the dependent variable (66 × 1) and X is the matrix of explanatory variables consisting of the intercept and the four factors (66 × 5). Since the residuals may exhibit autocorrelation and heteroskedasticity, standard errors were computed using the Newey–West method (HAC, lag = 3).

To test the hypotheses, three panel regression models were specified. Model A estimates the response of real construction output to real financing (H1):

$$\Delta ln\ Q_{it} = \alpha_i + \beta_1 \cdot \Delta ln(F_{it}/P_t) + \beta_2 \cdot \Delta ln(F_{i,t-1}/P_{t-1}) + \varepsilon_{it}, \quad (4)$$

where Q_{it} is real construction output in region i in year t; F_{it} is nominal financing of construction and installation works; P_t is the construction price index; α_i is the regional fixed effect. Model B explains the response of nominal construction output to financing and prices, and Model C explains the growth of regional construction cost (the implicit deflator) (H2 and H3):

$$\Delta ln\ Y_{it} = \alpha_i + \gamma_1 \cdot \Delta ln\ P_t + \gamma_2 \cdot \Delta ln\ F_{it} + \gamma_3 \cdot \Delta ln\ F_{i,t-1} + u_{it}, \qquad (5)$$

$$\Delta ln\ D_{it} = \alpha_i + \delta_1 \cdot \Delta ln\ P_t + \delta_2 \cdot \Delta ln\ F_{it} + \delta_3 \cdot \Delta ln\ F_{i,t-1} + v_{it}, \qquad (6)$$

where Y_{it} is nominal construction output and $D_{it} = Y_{it}/Q_{it}$ is the regional construction cost index. The logical link between Models B and C is important: since nominal output growth equals the sum of real output growth and cost growth, the share of the sum $\delta_2 + \delta_3$ in the sum of the coefficients $\gamma_2 + \gamma_3$ shows what part of additional financing is absorbed by price growth. In this study, this indicator is called the "price loss coefficient of financing".

The models were estimated with three estimators: pooled ordinary least squares (pooled OLS), fixed effects (FE) and random effects (RE). To select the estimator, an F-test for the significance of regional effects and the Hausman test were applied. The residuals were tested for cross-sectional correlation (Pesaran CD test), heteroskedasticity (modified Wald test for groupwise heteroskedasticity) and autocorrelation using the Wooldridge test. The stationarity of the variables was assessed with the Maddala–Wu (Fisher-ADF) panel unit root test.

3. Results and Discussion

Descriptive statistics for the variables of the first-stage model are presented in Table 2.

Table 2. Descriptive statistics of the multifactor model variables (January 2021 – June 2026, monthly, %)

Variable	Mean	Standard deviation	Minimum	Maximum	Correlation with ΔC
ΔC (price of construction and installation works)	0.724	0.664	–0.602	3.150	1.00
INF (consumer prices)	0.731	0.537	–0.300	2.762	–0.01
MP (material prices)	0.539	0.841	–0.904	4.306	0.92
W (wages, monthly equivalent)	1.201	0.662	–0.621	2.614	0.41
ER (import prices, 3-month average)	1.047	0.591	0.019	3.020	0.50

Note: *n = 66.*

According to the descriptive statistics, prices of construction and installation works increased by an average of 0.724 percent per month in 2021–2026 (about 9 percent in annual terms) [6]. The monthly change in material prices has the highest variability (standard deviation 0.841) and the highest correlation with prices of construction and installation works (r = 0.92). The relationships with import prices (r = 0.50) and wages (r = 0.41) are moderate, while the monthly correlation with consumer prices is practically zero (r = –0.01). Correlations among the explanatory variables are low ($|r| \leq 0.34$), and the variance inflation factors (VIF) range from 1.02 to 1.21 [7]. Hence, there is no multicollinearity problem (a VIF > 10 is usually considered a sign of a problem). The Augmented Dickey–Fuller test confirmed the stationarity of all monthly growth indicators. The estimation results are presented in Table 3.

Table 3. Estimation results of the multiple regression model of construction cost growth

Variable	Coefficient (b)	Standard error (HAC)	t-statistic	p-value	Beta	VIF	Significance
Intercept (β_0)	–0.0656	0.0935	–0.70	0.486	–	–	–
INF (β_1)	0.0669	0.0352	1.90	0.062	0.054	1.02	*
MP (β_2)	0.6457	0.0387	16.69	0.000	0.817	1.21	***
W (β_3)	0.1427	0.0457	3.13	0.003	0.142	1.12	***
ER (β_4)	0.2113	0.0670	3.15	0.003	0.188	1.15	***

Note: *n = 66; R^2 = 0.903; R^2 = 0.896; F = 141.38 (p < 0.001); s = 0.214; DW = 1.26. ***, * – significant at the 1% and 10% levels. Standard errors are computed by the Newey–West method (lag = 3). Beta – standardized coefficient.*

Substituting the estimated coefficients into model (2) yields the multiple regression equation of construction cost growth:

$$\Delta \ln \hat{C} = -0.066 + 0.067 \cdot INF + 0.646 \cdot MP + 0.143 \cdot W + 0.211 \cdot ER. \quad (7)$$

The goodness-of-fit indicators of the model were calculated as follows [8]. The residual sum of squares is RSS = $\Sigma(y - \hat{y})^2$ = 2.7939, the total sum of squares is TSS = $\Sigma(y - \bar{y})^2$ = 28.6956, and the explained sum of squares is ESS = TSS – RSS = 25.9017. Coefficient of determination: R^2 = 1 – RSS/TSS = 1 – 2.7939/28.6956 = 1 – 0.0974 = 0.9026. Adjusted coefficient of determination: R^2 = 1 – [RSS/(n – k – 1)]/[TSS/(n – 1)] = 1 – (2.7939/61)/(28.6956/65) = 1 – 0.04580/0.44147 = 0.8963. Thus, 90.3 percent of the monthly variation in prices of construction and installation works is explained by the four factors in the model. Fisher's criterion: F = (ESS/k)/(RSS/(n – k – 1)) = (25.9017/4)/(2.7939/61) = 6.4754/0.04580 = 141.38. This value is many times greater than the critical value (F_{cr}(4; 61; 0.05) = 2.52), i.e., the model as a whole is statistically significant. Standard error of the regression: s = √(RSS/(n – k – 1)) = √0.04580 = 0.214 percentage points [9].

The significance of each coefficient was tested with Student's t-test: t = b/SE(b). For material prices, t = 0.6457/0.0387 = 16.69; for wages, t = 0.1427/0.0457 = 3.13; for the exchange rate (import prices), t = 0.2113/0.0670 = 3.15. All three coefficients are significant at the 1 percent level (t_{cr}(61; 0.01) = 2.66) [10]. For inflation, t = 0.0669/0.0352 = 1.90, so the coefficient is significant at the 10 percent level (p = 0.062). In addition, the price homogeneity hypothesis was tested. If the prices of all resources rise by the same percentage, construction cost should rise by that same percentage, i.e., $\beta_1 + \beta_2 + \beta_3 + \beta_4 = 1$ [11]. The sum of the coefficients is 0.067 + 0.646 + 0.143 + 0.211 = 1.067. The Wald test does not reject this hypothesis (χ^2 = 0.57; p = 0.45). This confirms that the model is consistent with economic theory and that resource prices are fully passed through to construction cost.

The economic interpretation of the coefficients is as follows. A 1 percent increase in the price of construction materials, other things being equal, raises the cost of construction and installation works by 0.646 percent. This coefficient reflects the large share of materials in the estimated cost (according to the cost-estimate analysis in the author's other studies, materials account for 60–70 percent of direct costs). A 1 percent increase in import prices (the exchange rate) raises construction cost by 0.211 percent, and a 1 percent increase in wages raises it by 0.143 percent [12]. The direct effect of consumer inflation is small (0.067). This indicates that headline inflation is transmitted to construction cost mainly through materials, wages and import prices, i.e., indirectly. In terms of the standardized (beta) coefficients, the relative importance of the factors is as follows: material prices – 0.817, exchange rate – 0.188, wages – 0.142, inflation – 0.054 [13].

Based on the estimated equation, construction cost growth in 2021–2025 can be decomposed by factor. The annual contribution of each factor is calculated as $\beta_k \cdot \Sigma x_k$, where Σx_k is the sum of the factor's monthly growth rates over the year. For example, for 2025: the contribution of material prices is 0.6457 × 4.686 = 3.03 percentage points; the exchange rate, 0.2113 × 8.077 = 1.71 p.p.; wages, 0.1427 × 9.591 = 1.37 p.p.; inflation, 0.0669 × 7.067 = 0.47 p.p.; the intercept, –0.0656 × 12 = –0.79 p.p [14]. The growth predicted by the model is 5.79 percent, against actual growth of 5.48 percent (a difference of 0.31 p.p.). The results for all years are presented in Table 4 and Figure 1.

Table 4. Decomposition of the growth in the cost of construction and installation works by factor, percentage points

Indicator	2021	2022	2023	2024	2025	2021–2025	Share, %*
Actual growth (ΔlnC)	17.33	10.15	5.18	3.59	5.48	41.73	–
Material prices (MP)	10.07	2.95	1.08	1.73	3.03	18.87	41.3
Exchange rate (ER)	4.35	3.29	1.80	2.42	1.71	13.56	29.7
Wages (W)	2.98	2.41	2.04	1.45	1.37	10.25	22.4
Inflation (INF)	0.63	0.76	0.54	0.61	0.47	3.01	6.6
Intercept	–0.79	–0.79	–0.79	–0.79	–0.79	–3.93	–
Model-predicted growth	17.24	8.63	4.67	5.43	5.79	41.75	–
Residual	0.09	1.52	0.50	–1.84	–0.31	–0.03	–

Note: ** Relative to the sum of the contributions of the four factors.*

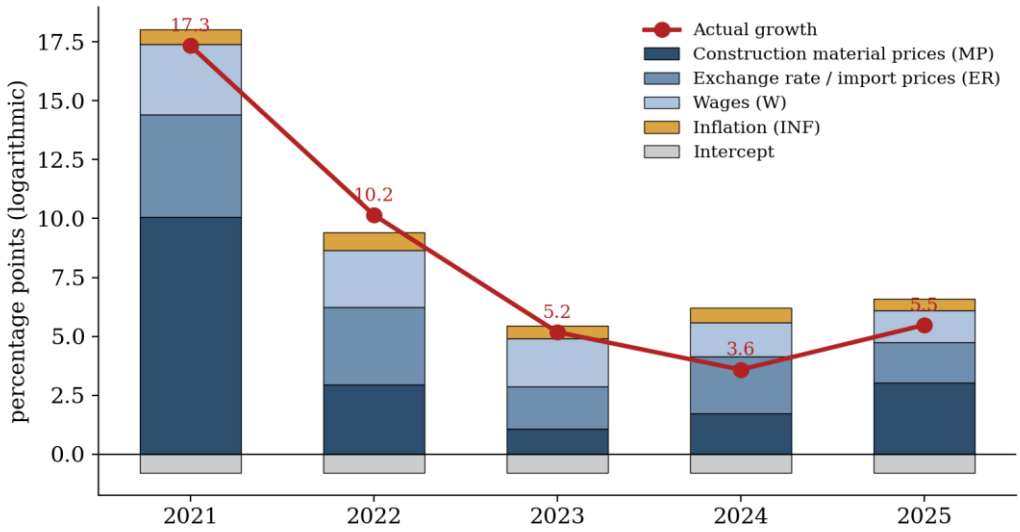


Figure 1. Decomposition of the growth in the cost of construction and installation works by factor, percentage points

According to the decomposition, the contributions of the factors to the total 41.7 percentage point (logarithmic) growth in the cost of construction and installation works in 2021–2025 were distributed as follows: material prices – 18.9 p.p. (41.3 percent of the sum of factor contributions), the exchange rate and import prices – 13.6 p.p. (29.7 percent), wages – 10.3 p.p. (22.4 percent), consumer inflation – 3.0 p.p. (6.6 percent) [15]. The construction cost growth in 2021 (17.3 percent) was mainly associated with a sharp increase in material prices (10.1 p.p.). This coincides with the post-pandemic surge in world prices for metals and other construction materials. In 2023–2025, by contrast, the contribution of materials declined, while the relative importance of wages and the exchange rate increased. This shows that managing construction cost requires taking into account not only material prices but also the labor market and currency risk [16].

The diagnostic results for the panel models are presented in Table 5.

Table 5. Diagnostic test results for the panel models

Test	Model A	Model B	Model C	Conclusion
F-test for regional effects (p)	0.983	0.995	0.365	Regional effects insignificant
Hausman test, χ^2 (p)	0.224 (0.894)	0.272 (0.965)	0.277 (0.964)	FE–RE difference insignificant
Pesaran CD test (p)	13.31 (0.000)	12.59 (0.000)	7.53 (0.000)	Cross-sectional correlation present
Modified Wald test, χ^2 (p)	69.5 (0.000)	61.5 (0.000)	83.0 (0.000)	Heteroskedasticity present
Wooldridge AR(1) test, t (p)	1.71 (0.087)	2.48 (0.013)	1.58 (0.115)	Autocorrelation in Model B
Fisher-ADF: dependent variable, χ^2	117.5***	88.8***	83.5***	Stationary
Fisher-ADF: financing, χ^2	131.0***	125.4***	125.4***	Stationary

Note: **** – the null hypothesis of a unit root is rejected at the 1% level.*

The diagnostic results show the following. The Fisher-ADF test rejects the null hypothesis of a unit root for all variables at the 1 percent significance level, i.e., the variables expressed as growth rates are stationary and there is no risk of spurious regression. The F-test for the significance of regional effects does not reject the null hypothesis in any of the three models, and the Hausman test indicates no systematic difference between the FE and RE estimates [17]. Hence, time-invariant differences between regions play no significant role in a model specified in growth rates, and all three estimators yield almost identical results. This is important evidence of the robustness of the results. At the same time, the Pesaran CD test indicates strong cross-sectional correlation in the residuals. This is natural, since all regions are exposed to common macroeconomic shocks (currency liberalization, the pandemic, government programs) [18]. Groupwise heteroskedasticity was also detected, and Model B shows signs of first-order autocorrelation. Therefore, standard errors were computed using the Driscoll–Kraay method, which is robust to cross-sectional correlation, heteroskedasticity and autocorrelation. This approach is recommended for macro panels with large N and moderate T. The estimation results are presented in Table 6.

Table 6. Estimation results of the panel regression models (Driscoll–Kraay standard errors)

Variables	A (FE)	A (RE)	B (FE)	B (RE)	C (FE)	C (RE)
Dependent variable	Δln Q	Δln Q	Δln Y	Δln Y	Δln D	Δln D
Δln(F/P)	0.152*** (0.024)	0.151*** (0.026)	–	–	–	–
Δln(F/P), lag	0.047 (0.030)	0.050* (0.030)	–	–	–	–
Δln P	–	–	0.735*** (0.156)	0.736*** (0.146)	0.382*** (0.047)	0.381*** (0.045)
Δln F	–	–	0.170*** (0.024)	0.169*** (0.024)	0.038*** (0.013)	0.038*** (0.012)
Δln F, lag	–	–	0.120*** (0.041)	0.125*** (0.040)	0.044*** (0.005)	0.045*** (0.005)

Constant	0.099*** (0.014)	0.099*** (0.013)	0.067** (0.028)	0.066** (0.027)	0.046*** (0.009)	0.045*** (0.009)
Number of observations	154	154	154	154	154	154
R^2 (within)	0.153	0.153	0.397	0.397	0.349	0.349
F-statistic	12.46	13.45	30.10	32.80	24.46	25.23

Note: *Standard errors in parentheses. ***, **, * – coefficient significant at the 1, 5 and 10 percent levels, respectively.*

The results of Model A confirm hypothesis H1. Current-year growth in real financing of construction and installation works has a statistically significant positive effect on real construction output: in the FE model, the elasticity is 0.152 (significant at the 1 percent level). The lagged effect is 0.047 but is only marginally significant with Driscoll–Kraay standard errors ($p = 0.12$; $p = 0.10$ in the POLS and RE models). The total (short- and medium-term) elasticity is 0.199, which means that a 10 percent increase in real financing raises real construction output in a region by about 2 percent. The hypothesis that the sum of the elasticities equals one is strongly rejected (Wald statistic 264.4; $p < 0.001$) [19]. Such a low elasticity of real output with respect to financing is explained by two factors. First, investment in construction and installation works is directed to projects lasting more than one year, whose effect is not fully reflected in the reporting year. Second, a large part of regional construction output (private housing construction, small enterprises) is financed from sources that are not fully captured by investment statistics. At the same time, the low elasticity also indicates that the regions have a limited capacity to absorb additional funds.

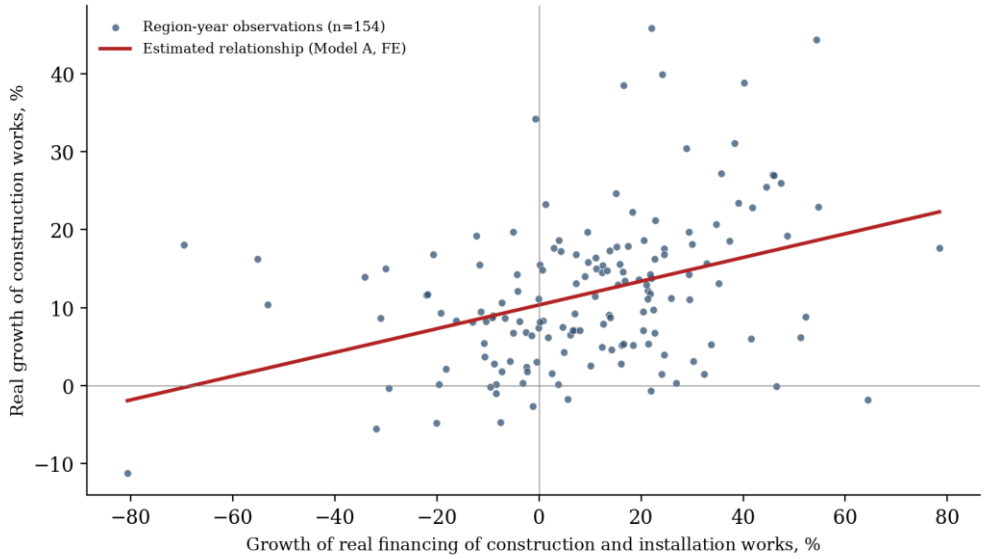


Figure 2. Relationship between growth in real financing and real construction output in the regions

Figure 2 shows a scatter plot of region-year observations and the relationship estimated with Model A. The plot shows that even when real financing contracted sharply (by up to 80 percent) or rose sharply (by up to 78 percent), growth in real construction output varied within a relatively narrow range (from –11 to +46 percent). This confirms the inertial nature of construction production and the partial transmission of financing fluctuations to real output.

The results of Models B and C confirm hypotheses H2 and H3. The coefficient of the national construction price index is 0.735 for regional nominal construction output and 0.382 for regional construction cost ($p < 0.001$ in both cases). This shows that the general increase in construction prices is transmitted to the regions not fully but at about

38 percent. The remainder is cushioned by the specific features of regional markets (local materials, wage levels) [20]. The most important result is the financing coefficients in Model C: growth in nominal financing of construction and installation works raises regional construction cost with an elasticity of 0.038 in the current year and an additional 0.044 in the following year (both coefficients are significant at the 1 percent level). In other words, a 10 percent increase in financing raises regional construction cost by a total of 0.8 percentage points over two years. This is direct econometric evidence of the existence of the demand channel.

Combining the results of Models B and C makes it possible to estimate how additional financing is divided between real output growth and price growth. The total elasticity of nominal construction output with respect to financing is 0.290 (0.170 + 0.120), of which 0.081 (0.038 + 0.044) corresponds to cost growth and the remaining 0.209 to real output growth [21].

Thus, the price loss coefficient of financing is 27.9 percent: about 28 percent of the nominal output growth generated by additional funds directed to construction and installation works is absorbed by construction cost growth rather than creating new real output. This result has important practical implications for centralized capital investment policy. A sharp increase in construction funding under government programs, especially where regional resource markets (construction materials, skilled labor, machinery) are constrained, loses part of its efficiency through price pressure [22]. Therefore, spreading program financing evenly over several years is more efficient than sharply increasing it within a single year. This is also closely related to the problem of uneven disbursement during the year noted in the author's other studies.

It is also instructive to compare the results with international studies. According to International Monetary Fund estimates, the average country loses about 30 percent of the potential gains from public investment due to inefficiencies in its management [23], [24]. B. Flyvbjerg and co-authors found that costs in infrastructure projects exceed plans by an average of 28 percent. The price loss coefficient of financing obtained in this study (27.9 percent) is of the same order of magnitude as these estimates, confirming that in Uzbekistan, too, construction cost growth is one of the main factors reducing the efficiency of public investment. At the same time, whereas the international estimates reflect losses at the level of individual projects, the coefficient calculated in this study represents a macroeconomic loss at the level of the regional construction market. The two approaches complement each other [25], [26].

Robustness checks confirm the main conclusions: the effect of financing on real construction output is positive and statistically significant in all panel specifications, ranging from 0.10 to 0.19. The lagged effect is also positive in all specifications [27]. This is explained by the fact that construction projects are financed over several years and work carries over into the following year. The low explanatory power of the national-level time-series model ($R^2 = 0.10$) and the insignificance of its coefficients confirm that the number of observations is insufficient. This, in turn, shows the need to base the main conclusions on the panel models.

4. Conclusions

The study leads to the following conclusions. First, the multifactor model based on monthly data for 2021–2026 explains 90.3 percent of the variation in prices of construction and installation works: $\Delta \ln \hat{C} = -0.066 + 0.067 \cdot INF + 0.646 \cdot MP + 0.143 \cdot W + 0.211 \cdot ER$. The coefficients of material prices, wages and import prices are significant at the 1 percent level, whereas consumer inflation is significant only at the 10 percent level. The sum of the coefficients is statistically equal to one, i.e., resource prices are fully passed through to construction cost. Second, in 2021–2025, materials accounted for 41.3 percent of construction cost growth, the exchange rate for 29.7 percent, wages for 22.4

percent, and headline inflation for only 6.6 percent. Third, panel models for 14 regions showed that the total elasticity of real construction output with respect to real financing is 0.199. Fourth, growth in financing raises construction cost through the demand channel: 27.9 percent of the nominal output growth generated by additional funds is absorbed by price growth. These results are robust across all alternative specifications. Based on the results, the following recommendations were developed: when indexing the cost of construction projects, use a formula composed of material, wage and import price components and calibrated on the model elasticities instead of the consumer price index; distribute financing evenly across years and quarters instead of sharply increasing program volumes; plan program limits in real terms and apply adjustment coefficients in regions with high price pressure; integrate the assessment and management of price risk in contracting enterprises into their corporate governance systems.

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