Clarification for "Not a Border Crisis, but a Labor Market Crisis: The Often Overlooked 'Pull' Factor of US Border Crossings"

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This note documents a minor issue in one statistic reported in the paper Bahar (2025) which despite not affecting the main findings of the paper, for transparency I offer a clarification.

In Bahar (2025) I reported an Engle and Granger (1987) cointegration test statistic of -18.829 obtained by regressing the first difference of *log_crossings* on the first difference of *log_openings* and testing the residuals for a unit root. In the strict mechanical sense, the statistic rejects the no-cointegration null at the 1% level. But presenting this test was an oversight, and an unnecessary one: Engle-Granger is designed for two $I(1)$ series, and as documented below, *log_openings* is $I(1)$ while *log_crossings* is closer to $I(0)$ once an endogenous structural break is admitted. Differencing then yields a regression of one $I(0)$ series on another, producing stationary residuals almost mechanically; the statistic accordingly has no useful economic interpretation and should be disregarded. Using appropriate methods, however, the substantive results remain the same, as I explain next.

Unit-root diagnostics give a clear $I(1)$ verdict for *log_openings* and a mixed verdict for *log_crossings*, which is closer to $I(0)$ once a 2020 structural break is admitted: a Zivot and Andrews (1992) test with an endogenous break locates the break at 2020m5 (May 2020, the COVID-19 disruption) and rejects the unit-root null with $t = -6.334$, well below the 1% critical value of -5.34. Therefore both mixed integration orders and an apparent structural break require cointegration tests other than EG ones. In fact, the Pesaran et al. (2001) ARDL bounds test, valid for any combination of $I(0)$ and $I(1)$ regressors (provided no variable is $I(2)$), rejects the "no cointegration" hypothesis at the 10% level in its primary specification (Column 1 of Table 1), but fails to reject once a deterministic linear trend is imposed (Column 2). The Gregory and Hansen (1996) test, the cointegration analog admitting an endogenous structural break, accommodates such a break in one of four ways: in the intercept alone (Model C), in the intercept and a deterministic trend (C/T), in the intercept and the slope coefficient on the regressor (C/S), or in all three (C/S/T). It rejects the hypothesis of "no cointegration even allowing for one break" at the 1% level in all four model variants (Columns 3 to 6 of Table 1), with estimated break dates varying across specifications and the trend-augmented models locating the break at 2020m9 and 2020m2 (the COVID-era disruption).

Table 1: Cointegration tests, monthly data

	ARDL bounds (t)		Gregory–Hansen (Z_t, break)			
	Primary	+Trend	C	C/T	C/S	C/S/T
Test statistic	-3.10^{*}	-2.95	-5.41^{***}	-7.93^{***}	-5.67^{***}	-7.93^{***}
Break date	—	—	2011m2	2020m9	2014m3	2020m2
N	282	282	286	286	286	286

Note: The table presents results of cointegration tests between *log_crossings* and *log_openings* on monthly data, sample 2001m4–2024m9. Columns 1 and 2 present ARDL bounds for primary specification and primary specification including deterministic trends, respectively (critical values from Kripfganz and Schneider, 2020 at 1%/5%/10%: for the primary specification $-3.82/-3.22/-2.91$ and for the primary+trend specification $-4.28/-3.69/-3.39$). Columns 3 to 6 report Gregory–Hansen test for level shift (C), level + trend (C/T), regime shift (C/S), and regime + trend (C/S/T), with critical values from Gregory and Hansen (1996, Table 1) at 1%/5%/10%: C $-5.13/-4.61/-4.34$, C/T $-5.45/-4.99/-4.72$, C/S $-5.47/-4.95/-4.68$, C/S/T $-6.02/-5.50/-5.24$. Break dates are estimated endogenously and reported in each of the GH models. Differences in N across columns reflect lag selection in ARDL ($N = 282$) relative to the full GH sample ($N = 286$).
*, **, *** denote rejection of the null at the 10%, 5%, 1% levels.

That said, the cointegration verdict is not load-bearing for the estimates, the interpretation, or the policy conclusions of the paper. The paper's main short-run regression, Specification B2, is

$$\Delta^{12} log\_crossings_t = \alpha + \beta\,\Delta^{12} log\_openings_t + \gamma_1\,\Delta^{12} log\_crossings_{t-1} + \gamma_2\,\Delta^{12} log\_openings_{t-1} + \theta\,\widehat{u}_{t-1} + \varepsilon_t,$$

where $\widehat{u}_{t-1}$ is the lagged residual from the levels regression of *log_crossings* on *log_openings*. In Bahar (2025) this term was described as an error-correction term (ECT), the natural reading if the series are cointegrated. Operationally, however, it enters B2 simply as a lagged residual control — absorbing variation in $log\_crossings_t$ not contemporaneously explained by $log\_openings_t$ — and B2 does not require the structural equilibrium reading. Consistently, Table 2 shows that the short-run elasticity β is essentially identical (0.463 vs. 0.464) with or without this control. B2 also fails the Granger and Newbold (1974) spurious-regression signature: $\Delta^{12} log\_crossings_t$ rejects the unit-root null (ADF $Z(t) = -2.69$, $p = 0.004$), B2 residuals reject it decisively ($Z(t) = -5.53$, $p < 0.001$), and the Durbin–Watson statistic is 1.35 — far from the near-zero values Granger and Newbold (1974) flag as the spurious-regression telltale. Whatever one concludes about cointegration, the paper's substantive results are unaffected.

Table 2: Equation B2 short-run elasticity, with and without the ECT control

Dependent variable: Δ^{12} log(border crossings)		
	(1) Published	(2) Without ECT
Job Openings Rate (logs)	0.463	0.464
	(0.056)***	(0.056)***
ECT ($\hat{u}_{t-1}$)	−0.010	—
	(0.024)	
N	273	273

The table reports OLS estimation of Specification B2 with and without the ECT control. Lagged controls ($\hat{\gamma}_1$ and $\hat{\gamma}_2$) are estimated but not displayed. Robust standard errors in parentheses.
$^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

References

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